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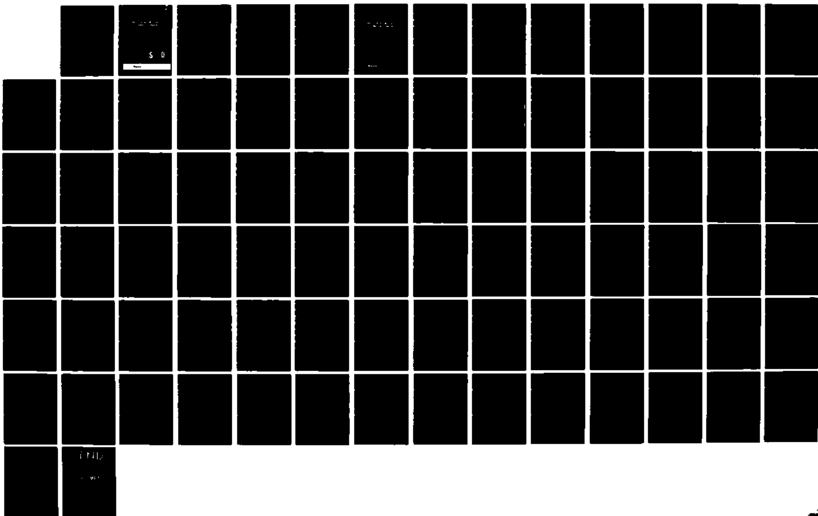

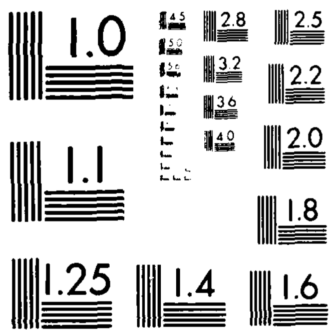


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# Reenlistment Bonuses and Retention Behavior

James R. Hosek, Christine E. Peterson







Rand

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This Report considers three outcomes for military personnel nearing a reenlistment decision point--reenlist, extend, or leave--and investigates whether bonuses increase the proportion of personnel staying in service (the retention rate), and whether bonuses affect the selection of longer terms of service (ie, raise the reenlistment rate over the extension rate). The authors used continuation rate data from the Defense Manpower Data Center for mid-FY76 through FY81. For each of over 500 occupations across the four services they computed reenlistment, extension, and retention rates at six month intervals, providing a total of 11 observations for each occupation in the analysis file. To these data they added information on reenlistment bonus coverage and amount, a military/civilian wage index variable, the unemployment rate, the percent of personnel without a high school degree or GED, and the percent black. Their analysis supports the contention that lump sum bonuses are more cost effective than installment bonuses in increasing the expected manyears of service in a military occupation.

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# Reenlistment Bonuses and Retention Behavior

James R. Hosek, Christine E. Peterson

March 1985

Prepared for the
Office of the Assistant Secretary of Defense/
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# PREFACE

Reenlistment bonuses constitute a small fraction of the annual budget for military compensation, but their role is an important one. By supplementing basic military compensation, they help prevent personnel shortages in occupations critical to the capability of the force. The inherent efficiency of reenlistment bonuses as a component of military compensation stems from their being selectively assignable: bonus dollars may be allocated to where they are most needed.

Still, several questions remain concerning the effects of bonuses on retention behavior. For example, if reenlistment bonuses increase the retention rate, to what extent do they encourage longer commitments and greater expected manyears of service? Do the retention effects of bonuses parallel those of military pay? Is it reasonable to expect that higher bonuses can offset the adverse retention effects of lower unemployment? And finally, are lump sum bonuses really more cost effective than installment bonuses?

This analysis originated with the issue of lump sum versus installment, yet the development of methodology and data suitable for that question permitted insight into the others. This research was prepared by Rand's Defense Manpower Research Center at the request of the Office of the Assistant Secretary of Defense, Manpower, Installations, and Logistics (Contract MDA903-83-C-0047).

# SUMMARY

How effective are reenlistment bonuses for improving the retention of enlisted personnel? Is it more cost effective to pay reenlistment bonuses in a lump sum or in installments? And how can reenlistment bonuses help counter the effects on retention of declining unemployment rates? When faced with questions of retention policy such as these, policymakers must consider not only what proportion of personnel are retained but also the average length of their new contracts. Changes in military pay, bonuses, and the unemployment rate that lead to the same change in the *retention rate* can nevertheless have different effects on *expected manyears*. Models of retention behavior not allowing for these differential effects may give policymakers a misleading impression of the consequences of such changes.

To assess effects on expected manyears, we distinguish three rather than two outcomes for personnel nearing a reenlistment decision point: not simply leave or reenlist, but leave, reenlist, or extend. The reenlistment rate refers to personnel who select a term of service over two and up to six years; the extension rate refers to personnel who select a term of service two years or less. The reenlistment rate plus the extension rate equal what we call the retention rate. This accounting framework permits the investigation of whether bonuses increase the proportion of personnel staying in service (the retention rate) and whether they affect the selection of longer terms of service (raise the reenlistment rate relative to the extension rate). Most previous bonus studies ignored extensions and dealt only with reenlistment behavior among the population of reenlistees and leavers.

The opportunity to examine the cost effectiveness of lump sum versus installment bonuses comes from a natural experiment: The method of bonus payment switched from installment to lump sum in April 1979. We utilize continuation rate data from the Defense Manpower Data Center for the period mid-FY76 through FY81. Individual military occupations form the basic unit of observation. For each of over 500 occupations across the four services we compute reenlistment, extension, and retention rates at six month intervals, providing a total of 11 observations for each occupation included in our analysis file. The rates are computed at both the first term and the second term reenlistment points. To these basic data we add information on reenlistment bonus coverage and amount, a military/civilian wage index variable, the unemployment rate, the percent of personnel without a

high school degree or GED (Certificate of General Educational Development), and the percent black.

Just as higher bonuses are expected to improve retention outcomes, a decline in retention outcomes can trigger greater bonus utilization. The simultaneity of this relationship can cause biases in the bonus coefficients as well as those of other variables. Going beyond previous analyses, we develop an econometric framework to control for simultaneity bias. The framework identifies and controls for two sources of bias, one from persistent yet unobserved factors affecting an occupation's reenlistment, extension, and retention rates, and the other from autocorrelation in these rates from one period to the next.

We find that reenlistment bonuses, whether lump sum or installment, increase the reenlistment and retention rate and decrease the extension rate. In other words, reenlistment bonuses increase not only the proportion of personnel who stay but also the proportion of stayers choosing longer terms of service. This pattern implies that higher reenlistment bonuses can increase the expected manyears of active duty service in an occupation.

We also find that lump sum bonuses are more cost effective than installment bonuses in increasing the expected manyears of service in a military occupation, at least at the first-term retention point. The added advantage of lump sum bonuses comes primarily from shifting personnel in an occupation from extension to reenlistment and secondarily from increasing the proportion of personnel who choose to stay in the occupation. (Evidence on the cost effectiveness of lump sum bonuses at the second term reenlistment point is inconclusive because of data limitations.)

The current method of bonus payment, begun in January 1982, blends the lump sum and installment approach. Half of the bonus is paid at the start of the reenlistment term, and the remainder is paid in annual installments. Precisely because the current method is a blend, it will be less cost effective than pure lump sum. We recommend that bonuses should be paid in lump sum.

Our findings imply that higher bonuses can counteract the effects of lower unemployment on retention and term length. The pattern of effects caused by a higher unemployment rate parallels that of higher reenlistment bonuses: higher reenlistment and retention rates, and lower extension rates.

Higher military wages can also offset lower unemployment but, unlike bonuses, would increase the proportion of stayers with short terms. Contrary to the result with higher bonuses, an increase in the military/civilian wage index *increases* the extension rate as well as the reenlistment rate and the overall retention rate. Moreover, the

absolute increase in the extension rate is roughly twice as large as the absolute increase in the reenlistment rate. Therefore, although expected manyears clearly increase, a higher fraction of those staying select short commitments. Given a reenlistment bonus increase and a military pay increase that each produces the same predicted increase in the retention rate for a military occupation, the bonus increase would be associated with a greater increase in expected manyears of service.

In most cases, our estimates of the first and second term effects of a variable were quite similar. This is true for the reenlistment, extension, and retention rate effects of installment bonuses, lump sum bonuses, the military/civilian wage index, and the percent black. However, for the unemployment rate, the second term effects were larger in absolute value than the first term effects, and for the percent without a high school diploma or GED the differences between first and second term were sporadic.

Controlling for simultaneity bias in the bonus coefficients proved to be crucial. Not doing so would have produced a lower first term bonus effect on the reenlistment rate and a *negative* second term bonus effect. The source of simultaneity bias associated with persistent yet unobserved factors associated with a military occupation had the greatest effect on our coefficients. In comparison, autocorrelation had little effect on the results; pooling observations across occupations appeared to eliminate any bias originating from autocorrelation. However, if we had sought to obtain bonus effects for individual occupations, or for small groups of occupations, autocorrelation might have been a problem.

On net, our analysis substantiates the view of reenlistment bonuses as a potent, versatile component of military compensation. Bonuses may be turned on or off rapidly and targeted on critical skills. They not only increase the retention rate but induce personnel to reenlist rather than extend, thereby increasing expected manyears. The reduction in extensions and increase in reenlistments probably gives personnel managers greater flexibility in planning rotation and new duty assignments, insofar as more personnel are committed to longer terms of service. Bonuses help alleviate transient, unexpected personnel shortages. Less widely recognized, bonuses also provide quasi-permanent pay differentials by occupation, to overcome persistent differences in work conditions and civilian earnings opportunities. Military occupations having a bonus in one period, we observe, tend to have a bonus in the next. Bonuses could be used to shape the mid-career force (years of service 4-12) in terms of both size and skill composition. Adding or increasing a bonus would promote retention in and retraining into areas where mid career personnel are most needed,

and presumably would also promote retraining into those areas from other, less critical areas. Such use of bonuses will not necessarily affect the number of personnel who continue on toward retirement eligibility. Other research suggests that those influenced by bonuses to stay an extra term are more likely to leave the service when that obligation is completed, other things equal. Finally, bonuses have the potential to offset the negative retention effects of improved national employment conditions, and to do so rapidly. That potential can be achieved, of course, only to the extent that bonus budgets permit a sufficient response.

The most immediate application of our analysis, beyond the question of lump sum versus installment bonus cost effectiveness, lies in forecasting reenlistment, extension, and retention rates for first and second term personnel. With some additional work, the consequences of changes in bonus coverage, bonus amount, and bonus allocation between first and second term reenlistment can be forecast at an aggregate level. Such forecasts can be integrated with research on trends in the personnel structure of the enlisted force.

This report concerns retention behavior at the DoD level; further work could provide estimates by service and occupational group. The dataset should be enhanced to permit analysis of reenlistment eligibility, promotion opportunity, and previous bonus payment, as well as to improve the demographic controls.

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Robert Brandewie of the Defense Manpower Data Center provided data on occupation-level continuation rates. Alice Mackey, Office of Enlisted Personnel Management, OASD(MI&L), supplied bonus histories for all occupations and answered many questions. Captain Lester Carl, Colonel Frederick Pang, Lieutenant Colonel Christopher Somers, and Colonel Harry H. Thie, also through the Office of Enlisted Personnel Management, offered support and encouragement throughout the project. Rand colleague Robert Bell guided us in weighting the data and handling the problem of simultaneity bias. Richard Fernandez, Richard Buddin, and Daniel Kohler gave us many useful comments, as did Glenn Gotz and Susan Hosek. We received excellent programming assistance from Betty Mansfield, Larry Painter, and Karl Schutz; and Susan Marquis and Michael Murray provided thoughtful, helpful reviews.

# CONTENTS

periods of data and accounted for 83 percent of the first term popula-tion. For the second term, 40 percent, or 542 occupations, had nonempty cells and accounted for 87 percent of the second term population. Table 2 shows the distribution of these occupations by service. We expect the deletion of occupations with empty cells to have little effect on our empirical estimates because our estimation procedure assigns small weights to small occupations.[4]

## DEFINITION OF VARIABLES

### Reenlistment, Extension, and Retention Rates

We define the reenlistment, extension, and retention rates with reference to the population at the beginning of a six-month interval. The population makes reenlistment, extension, or leave choices during the period, and by the end of the period those choosing to remain are grouped according to their new lengths of obligation. The DMDC data do not explicitly identify extensions or reenlistments. However, most reenlistments involve obligations of three to six years, and extensions usually last less than a year, with most extensions being no more than two years. With this in mind, we defined extenders as those with new obligations of 24 months or less, and reenlisters as those with new obligations of 25 to 72 months.[5] The extension rate, then, equals the

Table 2

NUMBER OF OCCUPATIONS IN ANALYSIS FILE

| Service | First Term | Second Term |
|---|---|---|
| Army | 213 | 210 |
| Navy | 84 | 87 |
| Marines | 152 | 109 |
| Air Force | 121 | 136 |
| DoD | 570 | 542 |

[4] In fact, the total continuation rates for each six-month period in our analysis file are quite similar to those produced by DMDC based on all skills.

[5] Although only those with new obligations of 36 months or more are eligible for bonuses, we had to include those with 25-35 months of obligation as reenlisters because the grouped data we received did not separate out the 25-35 month obligations from the 25-72 group. However, as most reenlistees opt for 36 months or more, there are probably only a few in the 25-35 month obligation category.

for male personnel with 6 to 9 years of service. The personnel all faced a decision to be made within the six-month interval—namely, to reenlist, extend, or leave. The counts were distributed by the length of new service obligation as recorded at the *end* of the interval, 24 months or less and greater than 24 months. The categories reflect decisions to continue in service. Personnel choosing to leave the service would not incur a new service obligation and would not appear in either of the categories.[3] In addition to the counts, each observation contained the percentage of individuals who were black, and the percentage with less than a high school education.

Because of the small size of many occupations, empty cells existed for some six-month periods. To study the effect of lump sum versus installment bonuses, we needed to have sufficient numbers of observations both before and after April 1979, when the change in method of payment occurred. In addition, for our econometric model we needed the capability to obtain good estimates of variable means by occupation, and the more time periods available the better the estimates. More time periods per skill also provide a better chance to estimate the effects of time-varying variables such as military/civilian pay and the unemployment rate. For these reasons, we decided to use only those specialties with nonempty cells in all time periods.

The criterion of continuity through time resulted in our keeping larger occupations; occupations with small numbers of people are more likely to have an empty cell in some period (i.e., no one fell within the six month "expiration of term of service" window). Our file contained 1366 first term occupations and 1361 second term occupations. For the first term, 43 percent, or 570 occupations, had nonempty cells in all 11

specified our econometric model to allow for unobserved yet persisting differences in the eligibility criteria across occupations (see Sec. V). Some recorded eligibility information may be misleading. For example, eligibility status may be set after a person has left the service, not necessarily before. Individuals who separate up to three months early for the purpose of attending college may be declared "ineligible." Also, in some cases ineligibility may be overturned by a waiver process.

A previous study of the effect of variable reenlistment bonuses on first term reenlistment (Enns, 1977) suggested that the inclusion of ineligibles did not result in different bonus effects than if they were excluded.

[3]Because the counts are based on snapshots taken six months apart within an occupation, our separations may include individuals who first extended for a couple of months and then left the service before the end of the six-month period. Thus, our extension rate represents the proportion remaining in the occupation at the end of six months who had not yet formally reenlisted or separated. It does not reflect the proportion who first extend at their reenlistment point.

Separations within the occupation may also include any individuals who cross-train and reenlist into a *different* occupation by the end of the six month period. Those people will not appear in the obligation categories of *either* occupation as they were not in either occupation for the entire six months.

# IV. SAMPLE AND VARIABLES

The historical pattern of reenlistment bonus payment methods provides a natural experiment for analyzing the effect of lump sum versus installment bonuses. This meant we could use existing historical retention files maintained by the Defense Manpower Data Center, which supplied us with retention rate data for each military occupation during the period FY76 to FY81.[1] We supplemented that data with specially constructed variables on the bonus amount and the military/civilian wage index. This section describes the DMDC data, the sample extracted, and the construction of the dependent and explanatory variables used in the empirical analysis.

## DEVELOPMENT OF THE ANALYSIS FILE

Individual military occupations are the basic unit of analysis. For FY76 through FY81, observations on all enlisted military occupational specialties [Primary Military Occupational Specialty (PMOS), Air Force Specialty Code (AFSC) and Navy Rating], over 1300 specialties in all, were taken at six-month intervals covering the first half and the last half of each fiscal year, starting with the last half of FY76. Observations were divided into first term and second term groupings. Each first term observation refers to the number of *male* personnel in the occupation with 3 to 5 years of service who were within six months of the end of their enlistment term, as counted at the beginning of the interval.[2] Second term observations contain the same kinds of counts

[1]During that period an additional change occurred in the bonus program. In 1981 the ceilings on reenlistment bonuses were raised from $12,000 to $16,000 for nonnuclear skills and from $15,000 to $20,000 for nuclear skills. Bonus ceilings dictate the maximum size reenlistment bonus an individual can receive. Individuals in skills offering a high bonus multiple (5 or 6, for example) might bump against the ceiling if they want to reenlist for more than, say, four years. But there would be no financial reward for doing so because the level of the bonus would not increase. The effect of bonus ceilings, therefore, tends to be seen in the length of commitment for which one reenlists rather than in the decision to reenlist, extend or leave. Given that, and the fact that the ceiling change occurred during the last two periods of our data, we do not believe that our ability to estimate the effect of lump sum payment versus installment payment is much hampered by the ceiling.

[2]With our data we could not judge whether all of the people in a skill at the beginning of a period were eligible to reenlist. Anyone likely to be ineligible because of a disciplinary problem or a physical disability was deleted at the outset. However, because of differences in eligibility requirements over time and across services, those ineligible for other reasons were not separately identified or deleted. As a partial [illegible] this we

new assignment, the person may have to remain in his current assignment and location, delaying the time until his next promotion to the extent that depends on experience and performance in the new assignment. Similarly, training opportunities might be withheld until the person commits to another term of service. The uncertainty about the future path of unemployment and the possible disruption of career progression suggest that the decision to extend for the sake of better timing one's departure may not pay off.

The risk can be avoided by reenlisting rather than extending. Because the perceived risk will be greater the higher the unemployment rate, there should be a tendency for the extension rate to decline and the reenlistment rate to rise as the unemployment rate rises. It is an empirical question whether this tendency dominates the role of extensions in mediating the timing of a person's departure from the service: here, higher unemployment would mean a higher extension rate (and a higher reenlistment rate).

Table 1 displays the hypotheses. The hypothesis that the effects of a lump sum bonus will be larger in absolute value than those of an installment bonus appears in the form of double-plus and double-minus signs.

Table 1

HYPOTHESES

| | Reenlistment Rate | Extension Rate | Retention Rate |
|---|---|---|---|
| Bonus amount | + | – | + |
| Lump sum payment | ++ | – – | ++ |
| Military/civilian wage index | + | + | + |
| Unemployment rate | + | – or + | + |

The effect of the wage index depends on whether enlisted personnel are satisfied with their current occupation. For those who are satisfied, an increase in the military/civilian wage index should cause either extension or reenlistment of some personnel who had planned to leave and reenlistment of some who had planned to extend. That is, the extension rate and the reenlistment rate should rise. For personnel not satisfied with their occupation, a higher wage index should shift some individuals from the leave category to the extend category, where retraining for a different military occupation would be undertaken. This would increase the extension rate. The extension rate might also increase if personnel were drawn from the reenlistment category into the extension category, but we believe this pathway is negligible because the higher wage index presumably does little to change relative earnings among military occupations.

## UNEMPLOYMENT RATE

We hypothesize that the unemployment rate will be positively related to the retention rate, because changes in the unemployment rate reflect changes in civilian employment opportunities. Because higher unemployment means a lower likelihood of finding an acceptable civilian job, judged in terms of wage, work conditions, training opportunities, and other relevant dimensions, a member of a service is more apt to remain in the service. The effect on the reenlistment and extension rates is less obvious. If an increase in the unemployment rate signals a brief aberration in the labor market, then the extension rate might rise while the reenlistment rate remained unchanged. The rise in the extension rate would reflect the use of the extension option by enlisted personnel to alter the timing of their departure from the military. Those planning to leave would use extensions to postpone their exit until better employment conditions prevailed. But two factors add complexity. First, poor employment conditions might last longer than the extension, and the speed and direction of change in the unemployment rate is difficult to predict. Second, an extension could disrupt the person's career progression. Rather than being rotated [illegible]

corresponding growth in the present value of expected military earnings [illegible] related to pay grade, subsequent years of service, and promotion speed. If an [illegible] military pay caused major increases in retention, personnel [illegible] would [illegible] motion rates to mitigate prospective surpluses of personnel in [illegible] ranks. [illegible] would lessen the size of the increase in present value, but [illegible] still remain [illegible] An increase in the civilian wage index broadly reflects [illegible] civilian w[illegible] experience, skill, and tenure groups, hence in the present [illegible] employment [illegible] civilian jobs.

prorated share of the bonus for the portion of the term not completed. In practice, the enforcement of this provision has not led to full recoupment, partly because of the limited degree of abuse and partly the positive cost of enforcement.[2] The person's chances of retaining the full amount of the lump sum reenlistment bonus are thus fairly high even if he does not complete his term. (Over 80 percent of first-term reenlistees complete their second term of service, and approximately 88 percent of second-term reenlistees complete their third term of service.) In comparison, recipients of installment bonuses do not receive the installments unless they remain in service. This means that the recipient bears all the bonus-related risk associated with the possibility of leaving early.

Another facet of risk comes from the functioning of the capital market. A lump sum bonus can be used to finance a purchase immediately, but an installment bonus cannot until one borrows on it. Generally, borrowing in anticipation of future income is difficult. One can instead borrow from savings or from other sources of liquidity (e.g., credit cards), but doing so may reduce liquidity to a suboptimal level. One then runs the risk that the lower liquidity will be inadequate to meet planned expenditures plus contingencies. In our view, both discounting and the difference in riskiness make an installment bonus worth less to a person than its nominal lump sum equivalent.

## MILITARY/CIVILIAN WAGE INDEX

The retention decision depends on expected military earnings versus expected civilian earnings. Apart from bonuses, we do not have occupation level data on military and civilian earnings and instead rely on a military/civilian wage index. An increase in the index implies that it is more lucrative to remain in military service; but because the gain in expected military earnings will be approximately equal across military occupations, one military occupation will not necessarily gain relative to another.[3]

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[2] According to a report by the General Accounting Office (1982), about 70 percent of so-called unearned bonus dollars are not recouped. This figure apparently includes both enlistment and reenlistment bonuses. A 1982 Defense Audit Service report, cited in the GAO report, estimates that up to $69 million in reenlistment bonuses were paid from FY78-82 to reenlistees who left the service before completing their terms. Roughly speaking, if the average term were four years and the average time of departure two years, then the unearned portion of these bonuses would amount to about $35 million, or 2 percent of the $1.8 billion projected outlays for reenlistment bonuses from FY78-82. Using the GAO estimate of 70 percent, about $25 million would not have been recouped.

[3] An increase in current military pay will increase military earnings over several years, the duration of the real increase varying with future increases and future inflation. The

# III. HYPOTHESES

This section presents hypotheses that are testable with our data. The hypotheses concern four variables—reenlistment bonus amount, method of bonus payment, military/civilian pay, and the unemployment rate—and their effects on the reenlistment rate, the extension rate, and the retention rate. A variable might have no effect on the retention rate even though it increases the extension rate and reduces the reenlistment rate, or vice versa.

## REENLISTMENT BONUS AMOUNT

We expect an increase in the size of the bonus to increase the retention rate, increase the reenlistment rate, and reduce the extension rate in a military occupation. The retention rate rises because a higher bonus raises the return to that occupation relative to other military occupations and relative to civilian employment. The reenlistment rate increases and the extension rate declines because to receive a reenlistment bonus, personnel must reenlist with at least three years' service obligation.[1] Moreover, the higher the bonus in an occupation the lower the gain from extending in order to retrain and transfer to another occupation.

## METHOD OF BONUS PAYMENT

We expect the effects of a lump sum bonus to be in the same direction but larger in absolute value than the effects of the installment bonus. In particular, lump sum and installment bonuses of equal face value should have different sized effects because the recipient will discount an installment bonus, paid in annual installments over the term of service. Its present value will therefore be less than the lump sum bonus.

The two methods of payment also differ somewhat in their riskiness. A lump sum bonus is paid in full at the beginning of the term. If the recipient does not complete his term, the government can collect the

[1] [illegible] who extend for three or more years are also eligible for [illegible] bonuses. [illegible], such people are counted [illegible] our reenlisters [illegible] reenlistment [illegible] having a new expiration of term of service (ETS) date [illegible] later than the ETS date occurring during the [illegible] period of ob[illegible].

in the civilian sector, personnel not yet ready to obligate to another term of military service. Such personnel are unsure about whether to continue in service, and attempts to classify them as reenlistees or leavers probably could not be done with much confidence. As a side point, we suspect that the option to extend works to the person's advantage by giving him more flexibility regarding the timing of his reenlistment decision, and to the service's advantage by reducing the likelihood that he will sign a reenlistment contract with only a marginal commitment to fulfill it.

reenlistments. Extensions, or new obligations of 24 months or less, are not included as a possible outcome.

In our opinion the usual approach of excluding extenders from the denominator in a calculation of the reenlistment rate precludes investigation of one of the choices personnel can make. What factors draw personnel into or out of the extender category? If bonuses are such a factor, their effect on extensions obviously cannot be studied with extenders removed from the population. Moreover, the deletion of extenders in computing the reenlistment rate makes interpreting a change in that rate more difficult. For a fixed pool of individuals who in general may reenlist, extend, or leave, an increase in the conventional reenlistment rate (which excludes extenders from the denominator) can be caused either by an increase in the number of reenlistments (holding extensions constant) *or* by an increase in the number of extensions (holding reenlistments constant). Either of these increases will mean a decrease in the number of separations, so the conventional reenlistment rate must rise. But these separate types of increases in the reenlistment rate may not be equivalent from a policy perspective: An increase due to more reenlistments means more personnel have obligated for longer terms of service, while an increase due solely to more extensions means *no additional* personnel have obligated for longer terms.

Our accounting framework avoids such ambiguity. The base population includes all eligible personnel within a given time of the end of their current term of service. We define the *reenlistment rate* as the number of reenlistments over the base population, and the *extension rate* as the number of extensions over the base population. We also define a *retention rate*, equal to the sum of the reenlistment rate and the extension rate. By implication, the separation rate is one minus the retention rate.

Some may argue that including a category for extensions in our accounting framework is inappropriate because extenders have already committed to a stay or leave decision but their decision is not yet apparent in the current period's information.[1] Actually, extenders are a diverse group. Extensions help smooth the transition from one military occupation to another. Training for a new occupation may have to be done on extension as a precondition for entry into the new occupation. Extensions also provide leeway for personnel considering jobs

[1]If the intentions of the personnel were known and were accurate indicators of behavior, a case could be made for including extenders and classifying them as reenlistees or separations. But intentions are not recorded, and the practice of computing the reenlistment rate has been to wait to include the extenders until their reenlist or leave decision has been realized. This approach mixes people who are making a reenlist or leave decision at time $t$ with people who made their decision possibly six months or a year earlier, when pay, unemployment, and policy conditions might have been different.

# II. ACCOUNTING FRAMEWORK

Analyses of bonus effects can produce potentially differing results depending on the way the data have been organized. We have chosen a particular organization, or accounting framework, which follows *all* personnel who approach a reenlistment decision point. This enables us to analyze what proportion of personnel remain in service and how that proportion divides into short and longer term commitments. Our framework conforms to the manner in which the Defense Manpower Data Center tabulates personnel flows from one time period to the next. However, many previous studies of bonuses and reenlistment behavior have used different approaches.

A description of the choices personnel face near the end of a term of service will help explain the difference in approaches. These personnel, if eligible, can choose to stay in service or to leave. If ineligible they must leave. Those choosing to stay may sign a new enlistment contract (reenlist) or increase the length of their current contract (extend). Reenlistment contracts specify the length of the service obligation, which ranges from two to six years but is typically three or four. Contract extensions, by comparison, frequently last less than a year and generally do not exceed 24 months. Extensions are probably less well known than reenlistments, but they are not negligible. In our sample of personnel within six months of the end of their term of service, about one-third of the stayers at the first term reenlistment decision point are extenders (defined as those with new obligations of 24 months or less), as are about one-fifth at the second term reenlistment decision point.

Few studies of retention behavior have looked at extenders. Goldberg and Warner (1982) examined the determinants of reenlistment and extension rates in the Navy; Zulli (1982) looked at income effects on the decision to reenlist vs. extend *given* recommitment (if the individual has decided to stay, what is the probability he will reenlist instead of extend) for Navy personnel with 11–14 years of service. Cylke et al. (1982) included extenders among those they defined as reenlisters and therefore were estimating the effects of bonuses on retention. Other previous studies on the effects of bonuses, however, have excluded extenders from the analysis (e.g., Enns, 1977; Chow and Polich, 1980). These studies use the conventional definition of the reenlistment rate—namely, the number of reenlistments (new obligation of 25 months or more) divided by the number of separations plus

an obligation (new or old) of at least three years of service, but the size of the effect had not been quantified in earlier studies. We are able to quantify the effect in part because of a different accounting framework than previous analyses.

Because bonuses influence the choice between longer and shorter obligations, they affect the number of expected manyears available from a given cohort. We examine the change in expected manyears with respect to a change in bonuses, both lump sum and installment. Our results indicate that bonuses can lead to higher numbers of expected manyears and that lump sum bonuses produce greater expected manyears than installment bonuses of equal face value. That higher number of expected manyears appears to be due more to the shift from extension to reenlistment (from shorter to longer term commitments) than to retention of those who would otherwise have left active duty.

A secondary but still important purpose of the analysis is to compare the effects of reenlistment bonuses with those of pay and unemployment, which along with other variables must in any case be controlled in the empirical work to produce proper estimates of the bonus effects. Our methodology explicitly considers the question of simultaneity bias--just as higher bonuses affect retention, lower retention may lead to higher bonuses—and provides direct control for pay, unemployment, and demographic composition, and indirect control for such factors as work conditions and average promotion opportunity. We may then ask whether higher pay, like bonuses, induces stayers to select longer terms of service, and similarly whether higher unemployment does so. We can also approximate the size of the military pay or reenlistment bonus increase that would be required, on average, to preserve retention rates at a given level in the face of a decline in the unemployment rate.

# I. INTRODUCTION

The chief purpose of the study is to determine whether lump sum reenlistment bonuses are more cost effective than installment bonuses. Under the lump sum method of bonus payment, all of the bonus is paid when the new term of service begins. Under the installment method of payment, the bonus is paid in equal annual installments over the duration of the new term of service. Which of these methods of payment should be preferred depends on whether, for a given reenlistment bonus budget, one method leads to greater retention of personnel than the other.

To resolve this question, one would ideally undertake a controlled experiment in which military personnel were randomly assigned to receive either lump sum or installment bonuses. Because no experiment has been conducted, we use a historical change in the method of bonus payment, which switched from installment to lump sum in April 1979. Our empirical work, utilizing data from FY76 through FY81, seeks to determine if this change caused an improvement in retention and if so whether the improvement was great enough to justify the higher cost of lump sum than installment bonuses.

The analysis implies that at the first term reenlistment point, lump sum bonuses are more cost effective than installment bonuses. This finding holds over a plausible range of assumptions required to place installment and lump sum bonuses on an equal cost basis. The results for the second term reenlistment rate are inconclusive, which we attribute to the major shift in the use of second term bonuses occurring in the last years of the sample and thereby confounding our ability to estimate the lump sum effect. We made no attempt to estimate effects for third term bonuses because their use was not authorized until fiscal 1981 and so they were paid only in lump sums during the period covered by our data.

In addition to estimating the effect of lump sum versus installment bonuses, we obtain estimates of the direct effect of each kind of bonus. From these estimates it is possible to predict the change in the reenlistment rate that would result from a given change in bonus amount for either method of payment. Moreover, we investigate the extent to which bonuses induce personnel to select a longer service commitment in their occupation. Bonuses not only induce more people to remain in their occupation, we find, but cause those who remain to choose longer terms. This should be expected because reenlistment bonuses require

## FIGURES



## TABLES

Appendix

number of extenders at the end of the interval divided by the population at the beginning of the interval. The reenlistment rate equals the number of reenlisters divided by the same population, and the retention rate is the sum of the reenlistment and extension rates.[6]

## Bonus Variables: Presence and Amount

Two basic selective reenlistment bonuses were created: an indicator for a bonus being offered in the occupation and a representative amount of the bonus. Information on the bonus history by occupation told us in which periods a bonus was available and at what bonus multiple. From this we defined a variable indicating the *presence* of a bonus that equals "1" in periods when a bonus is present and "0" otherwise.[7] Using this variable, we tabulated the proportion of *occupations* in each period that offered a reenlistment bonus. The tabulations, based on our analysis file, appear in Table 3. As shown, the percent of occupations with first term reenlistment bonuses rose by a third from FY76 to FY81, from 30 percent to 40 percent. Second term bonus coverage rose far more rapidly, burgeoning from 8 percent to 43 percent, with the greatest growth occurring after FY79. The pattern of change in the incidence of second term bonuses raised doubt that the differential effects of lump sum versus installment bonuses could be estimated accurately, as the expansion occurred at the same time as the shift to lump sum payment.

For reference, Table 3 also shows the percent of *personnel* six months or less from the end of their term of service who are in bonus

[6] These rates are lower than if we had used a 12-month interval instead of a six month one. The six month window misses personnel who reenlisted or extended seven to 12 months before the end of their term of service. We do not have information on the number of such early reenlistments and extensions. However, the six month interval provides twice as many observations over our period of analysis and increases the potential for precise estimation of bonus, pay, and unemployment effects. This increase in precision may well offset the cost of missing some personnel, particularly if the retention behavior of the latter is not appreciably different from that of the personnel in the six month window. The question of behavioral differences probably cannot be settled until a longer series of 12-month observations becomes available.

[7] In regressions not reported here, we also included variables representing bonus timing. Indicators for when a change in the bonus pattern for an occupation occurred in the next period were created. These indicators were for (1) if a bonus would be added, (2) if the bonus would be removed, (3) if the bonus multiple would be increased, and (4) if the multiple would be decreased. Results suggested that some "gaming" of the bonus system may affect the level of extensions. Extension rates were higher if a bonus was to be added in the next period and lower if the bonus would be removed or decreased. The effects on reenlistment rates, however, were mixed: decreases in next period's bonus did increase reenlistment rates, but so did increases. Superior estimates should be possible as the number of periods of observation grows and the number of bonus changes within an occupation increases.

Table 3

BONUS INCIDENCE
(Percent)

| | Occupations | | Personnel | |
|---|---|---|---|---|
| Period | First Term | Second Term | First Term | Second Term |
| FY76 II | 30 | 8 | 29 | 11 |
| FY77 I | 27 | 11 | 29 | 14 |
| II | 29 | 14 | 35 | 18 |
| FY78 I | 34 | 16 | 38 | 18 |
| II | 34 | 16 | 39 | 20 |
| FY79 I | 36 | 21 | 42 | 20 |
| II | 36 | 22 | 43 | 23 |
| FY80 I | 36 | 28 | 36 | 23 |
| II | 36 | 28 | 33 | 24 |
| FY81 I | 40 | 42 | 37 | 42 |
| II | 40 | 43 | 35 | 42 |

occupations. Since mid-FY77 over a third of personnel at the first term reenlistment point were in occupations covered by a bonus. This proportion has remained fairly stable in comparison with the bonus coverage among personnel at the second term reenlistment point, whose coverage rose so dramatically.

Using the bonus multiple, we computed the *amount* of a typical bonus for each bonus skill. An individual's bonus amount equals his monthly base pay times the number of years he reenlists, times the bonus multiple for his occupation.[8] During our period of analysis, bonus multiples were integers ranging from one to six. Having grouped data, we needed a measure of monthly base pay typical of the group, thus giving the same base pay to all occupations. We used the base pay for paygrade E-4 with four years of service for first termers, and paygrade E-6 with seven or eight years of service for second termers.

[8] Our bonus variables do not include the "save pay" provisions for regular reenlistment bonuses (RRBs). Among our first term observations, RRBs were phased out by June 1978; among second termers RRBs were paid through the entire period but the proportion eligible began falling after June 1980. Because we had grouped data, we could only estimate the proportion eligible for an RRB based on the Years of Service (YOS) distribution in the skill at time $t$. Such a variable was created to test whether there were any possible RRB effects, but this variable proved to have small effects and was statistically insignificant. It was not retained in our final regression specification.

The term of reenlistment was assumed to be four years, although a three year term would have produced essentially the same empirical results. To compare across years, we converted bonuses into constant 1976 dollars.[9] Appendix G presents the values per bonus multiple in each period. These values were multiplied by the occupation's bonus multiple to obtain the bonus amount for the occupation in a given period.

### Method of Bonus Payment

To measure the effect of the change from installment to lump sum bonuses, the bonus presence and amount variables were interacted with an indicator for whether the time period was after April 1, 1979, when bonuses began to be paid as lump sums. The coefficients on these interacted variables will indicate the extent of the *difference* in the effects of a lump sum bonus relative to those of an installment bonus, after controlling for all other factors. With this information we compute the effects of installment and lump sum bonuses and the difference in those effects.

### Military/Civilian Wage

A military/civilian wage index was constructed by comparing the growth in military pay with the growth in an index of civilian pay, the average hourly wage in manufacturing.[10] Because military pay increases occurred only on an annual basis, we assigned half of the increase to the first half of the fiscal year and the rest to the second half. Having made these assignments, we reasoned that service personnel would be interested not only in the current level of military pay but also in the expected future level. The course of military pay is generally difficult

[9] This adjustment, however, does not control for the erosion of pay over the life of the installment bonus that might be caused by differences in inflationary expectations over the installment bonus era. If expectations differed markedly, then, the real value of an installment bonus would also differ, even though its nominal value in 1976 dollars were the same. However, during the period in which installment bonuses were being paid, inflation was low and constant compared with the lump sum period. Thus individuals may have had stable inflationary expectations, and we would not expect the real value of the installment bonus variable to have varied much if at all. Therefore, our estimates of installment bonus effects should not be affected by the possibility of changes in inflationary expectations. Of course, this is not a problem for lump sum bonuses because they are paid immediately. Converting lump sum bonuses to 1976 dollars is sufficient to keep the real value of a lump sum bonus dollar constant.

[10] Quarterly figures for average hourly manufacturing wage from the Bureau of Labor Statistics were used to represent civilian wages. The six month period ending in March received the wage from the first quarter of that calendar year; periods ending in September received the third quarter value.

to predict, but in practice the approximate size of the pay increase for the coming fiscal year is known by the middle of the current fiscal year. Those facing a reenlistment decision presumably take this into account. Therefore, the growth in military pay as perceived by the person was defined as the average of the increase assigned to the current six month period and the increase assigned to the next.[11] This procedure also has the advantage of smoothing the military pay series, thereby preventing the sawtooth pattern that occurs if only the current period's increase is used. The same kind of smoothing is done for the civilian wage series, the argument again being that people are interested in and can forecast near term changes in the civilian wage. For both our smoothed military and civilian pay series we created indexes set equal to 100 in the first half of FY76. These were used to form the ratio of military to civilian wage growth during the 11 periods covered by our data. The ratio was multiplied by 100 to convert it to an index of the military/civilian wage. Appendix H presents the smoothed growth rates, the indexes based on the smoothed rates, and military/civilian wage index.

## Unemployment Rate

We used the unemployment rate for the total civilian labor force as reported by the Bureau of Labor Statistics. Like the civilian wage, the unemployment rate assigned to a period was the quarterly average for the quarter in the last half of the period. However, preliminary empirical work indicated that the unemployment rate lagged one period performed better in the regression analysis than either the current rate or the rate lagged two periods. Given the way we assigned the unemployment rate to periods, the lagged rate was simply the unemployment rate for the quarter immediately preceding the current period.[12]

[11]Suppose, for example, the military pay rose 3 percent in one year and 4 percent the next. Start by assigning a 1.5 percent increase to each half of the first year and a 2.0 percent increase to each half of the second year. The perceived military pay increase in the second half of the first year would then be (1.5 + 2.0) divided by 2.

[12]Both the military/civilian wage index and the unemployment rate vary only across time and not across occupations. This makes these variables sensitive to time period indicator. To try and control for the effect on retention of the Congressionally mandated end strength ceiling, we included in earlier specifications an indicator for whether the period was the last six months of the fiscal year. If retention (or accessions) has been running above normal for the first half of the year, the services may have to restrict the number of reenlisters and extenders in the second half so as not to exceed the end strength ceiling; therefore, reenlistments and extensions may be lower in the last half of the year than in the first half. Our results upheld this hypothesis; however, the effect of the end strength indicator on the wage index and especially on the unemployment rate made it difficult to sort out the effects of the latter two variables. We [illegible] chose to present a specification without the end strength indicator.

## Demographic Variables

We defined two demographic variables, the percentage of males without a high school diploma or its equivalent, and the percentage of blacks. The high school nongraduate percentage reflects educational attainment at the reenlistment point. These variables control for compositional differences across occupations that could otherwise confound bonus, pay, and unemployment effects. For instance, if military/civilian wage tended to increase at the same time the percent blacks increased, and if blacks have systematically lower civilian opportunity wages than whites, then not controlling for the percent black would bias the wage effect on the dependent variables upward.

# V. ECONOMETRIC MODEL AND ESTIMATION PROCEDURE

A chief concern in specifying the econometric model is to obtain consistent estimates of the effects of reenlistment bonuses on the dependent variables. The possibility of bias exists because of the mutual causation between bonuses and the dependent variables, especially the reenlistment rate. Higher bonuses should increase the reenlistment rate, but a higher reenlistment rate should reduce the size of the bonus. If uncorrected, this reverse causality can be expected to impart a negative bias to the effect of the bonus on the reenlistment rate, making bonuses appear less effective than they are. Both lump sum and installment bonus effects would be underestimated, so the estimate of the difference between the two effects would probably be too small as well. In that case, the cost effectiveness of lump sum bonuses would subsequently be underestimated.

In describing the econometric model, we focus on the reenlistment rate equation because here the causality and institutional arguments seem most direct. However, similar arguments apply to that the extension rate and the retention rate equations, and in conducting the empirical analysis we apply the same econometric approach to all three equations. After discussing the econometric model, we discuss procedures followed in estimating the model.

## ECONOMETRIC MODEL

### Reenlistment Rate Equation

Assume that the reenlistment rate for a given military occupation depends on the bonus variable, the other explanatory variables, an occupation-specific intercept, and an error term. For each occupation we have $T$ observations, and there are $I$ occupations in all. For discussion purposes, let the form of the *reenlistment rate equation* be

$$r_{it} = \beta_1 B_{it} + \beta_2 X_{it} + \delta_i + \varepsilon_{it} \tag{1}$$

This specification constrains the coefficients for the bonus and the other explanatory variables to be equal across occupations. The constraint is necessary because we do not have enough observations to

estimate equations for each occupation but must pool observations across occupations.

Each occupation has its own intercept, which is represented by $\delta_i$. The intercept accounts for the net effect of *unobserved* factors that persist throughout the period of analysis. These "permanent" factors include the unchanging aspects of work conditions, rotation policy, promotion policy, career development opportunities, pattern of hours of work, reenlistment eligibility conditions, and the like. They also include the unchanging aspects of civilian sector jobs likely to be sought by personnel in this occupation. These aspects include wages, hours of work, fringe benefits, locations, and possibly others. By capturing the permanent effect of the unobserved factors, the intercept permits the occupation to have a persistently higher or lower reenlistment rate than the average.

## The Error Structure

The error term in the reenlistment rate equation has two components, one occupation-specific autoregressive, the other a simple time-varying component:

$$\omega_{it} = \nu_{it} + \epsilon_t, \text{ where } \nu_{it} = \rho_i \nu_{it-1} + \xi_{it}. \qquad (2)$$

The autoregressive error can arise because of changes in the unobserved factors listed above. Such changes might affect the reenlistment rate at the time they occur as well as future reenlistment rates. The future effects are assumed to be dampened. For example, there could be a change in the wages or employment conditions of relevant civilian jobs. This phenomenon might not be fully captured by the pay and unemployment variables included on our list of explanatory variables, but could nevertheless affect the reenlistment rate. A tightening or loosening of reenlistment eligibility criteria could also produce autoregressive errors, as could a change in the probability of promotion. Note that the autoregressive error links a component ($\nu_{it}$) of the occupation's error term from period to period. We assume that $\nu_{it}$ is uncorrelated with the other components of the error term for this occupation or with other occupations' error terms, and that the autoregression process is first-order.

Movements in the unobserved variables need not be in concert and might counteract one another's effect on the occupation's reenlistment rate. For instance, if promotion opportunities improved at the same time that civilian employment conditions worsened, then the effects would be reenforcing and would presumably produce a positive correlation from period to period ($\rho_i > 0$). Alternatively, suppose

reenlistment eligibility conditions were tightened in the second half of a fiscal year if the reenlistment rate in the first half had been high (relative to the level necessary to meet current manpower requirements); then the eligibility conditions were relaxed at the beginning of the next fiscal year to increase the chances of meeting the manpower requirements for that year. This pattern of behavior could produce a negative correlation ($\rho_i$).

The other two terms in the error are $\xi_{it}$ and $\epsilon_t$. The first of these accounts for randomness in the observed reenlistment rate and is *uncorrelated with the other components of the occupation's error term*, either in the current period or other periods, and is uncorrelated with the error terms of the other occupations. The term $\epsilon_t$, in contrast, is by definition the same for all occupations in a given cross section. The kind of unobserved factors giving rise to this term would include transitory changes in the national security posture caused by threat of conflict, or the possibility of alteration in the military compensation package, which would affect all personnel, not just those in a single occupation. For example, a rumor that retirement benefits would change in a certain way might affect reenlistment behavior across the board.

Having offered some motivation for the error structure, we summarize it specifically as Eq. (2) above plus the covariance matrix:

$$\Sigma \quad \sigma_\nu^2 \begin{bmatrix} P_1 & & & & & 0 \\ & P_2 & & & & \\ & & \cdot & & & \\ & & & \cdot & & \\ & & & & \ddots & \\ 0 & & & & & P_1 \end{bmatrix} + \sigma_\epsilon^2 \begin{bmatrix} I & I & \cdots & \cdots & I \\ I & I & & & \cdot \\ \cdot & \cdot & & & \cdot \\ \cdot & & \cdot & & \cdot \\ \cdot & & & \cdot & \cdot \\ I & \cdots & \cdots & \cdots & I \end{bmatrix} + \sigma_\xi^2 \begin{bmatrix} I & & & & & 0 \\ & I & & & & \\ & & \cdot & & & \\ & & & \cdot & & \\ & & & & \ddots & \\ 0 & & & & & I \end{bmatrix} \qquad (3)$$

where the submatrices $P_t$ and $I$ have the dimension $T \times T$. The submatrices $P_i$, $i = 1, \ldots, I$, result from the correlation between the $\nu$s for an occupation in different periods. The correlation between period $t$ and period $t + j$ is $\rho_i^j$; this correlation may differ from occupation to occupation because the $\rho_i$ are permitted to differ across occupations. However, the $\nu$s are not correlated across occupations; therefore, the matrix contains zeroes off the block-diagonal. The second matrix, a

square array of identity submatrixes ($I$), reflects the error variance generated by $\epsilon_t$ within and across occupations in a given period. (Recall that the term $\epsilon_t$ is the same for every occupation in a particular period.) The third matrix, containing block-diagonal identity matrixes, characterizes the variance from the "random noise" component $\xi_{it}$, which is uncorrelated across occupations and time periods. All off-diagonal elements are therefore zero.

## Bonus Determination Equation and Simultaneity

Equation (4) describes the factors determining the presence and level of a reenlistment bonus in an occupation. From DoD policy we know that bonus-setting depends on the extent of shortage in an occupation, the cost of procuring additional personnel to fill the shortage, and whether the occupation is judged to be "critical," a catchall term for the perceived importance of meeting the manpower requirements for the occupation. This implies a bonus determination equation such as:

$$B_{it} = \alpha_0 + \alpha_1(r_{it}{}^* - r^o_{it}) + \alpha_2 Z_{it} + \eta_{it} \tag{4}$$

where $r_{it}{}^*$ = service's desired reenlistment rate in occupation $i$ and period $t$, $r^o_{it}$ = expected reenlistment rate without a bonus, $Z_{it}$ = cost of additional personnel and criticality, and $\eta_{it}$ = error term. We use this equation for notional purposes and in fact lack the data to estimate it.

The term $r_{it}{}^*$ expresses the rate of reenlistment necessary to meet current manpower requirements. If manpower requirements were constant through time, if compensation and promotion policy were constant, and if the pool of eligibles to reenlist were constant, the $r_{it}{}^*$ would also be constant. But changes in weapons systems, deployment patterns, compensation, accessions, or attrition can change $r_{it}{}^*$ over time.

The term $r^o_{it}$ represents the reenlistment rate that would occur in the absence of a bonus in the occupation. Assuming the coefficients of the reenlistment rate equation were known, the no-bonus reenlistment rate forecast for period $t$ would be

$$r^o_{it} = \beta_2 X_{it} + \delta_i + \rho_i \mu_{it-1} \tag{5}$$

where $t - 1$ is the last period in which the reenlistment rate has been observed and $\mu_{it}$ is the difference between the outcome and the prediction for that period. (If the last observed outcome were not period $t - 1$ but, say, $t - j$, then $\rho$ would be raised to the $j$th power and the appropriate residual would be $\mu_{it-j}$.)

Substituting Eq. (5) into Eq. (4) makes explicit the sources of simultaneity bias in the reenlistment rate equation. The substitution gives

$$B_{it} = \alpha_0 - \alpha_1 r_{it}^* - \alpha_1(\beta_2 X_{it} + \delta_i + \rho_i \mu_{it-1}) + \alpha_2 Z_{it} + \eta_{it}. \quad (6)$$

The relationship shows that the bonus depends directly on the occupation intercept $\delta_i$ and the intertemporal correlation $\rho_i$. The higher either of these, the lower the bonus. Since $\rho_i$ is part of the error term in the reenlistment rate equation, we can expect it to be correlated with the bonus variable. This violates the assumption that the error term is distributed independently of the explanatory variables, which is necessary to obtain consistent estimates of the bonus (and other) effects. To restore the credibility of the independence assumption, we employ a procedure for removing autocorrelation from the error term, as explained below.

We have mentioned that the bonus also depends on the occupation intercept $\delta_i$. This would not seem to present a problem, for as the model has been presented the intercept is simply another explanatory variable. However, our database contains over 500 occupations, and with each requiring its own intercept the estimation of the model becomes impractical. One alternative is to view $\delta_i$ as being unobserved and hence as a permanent component in the error term. But if that approach were followed, the error term would again be correlated with the bonus variable and a simultaneity bias would result. To eliminate the bias, the permanent component would have to be purged from the error term. A second alternative is to difference the observations for each occupation from period to period. Although this eliminates the intercept, it also could impart autocorrelation to the differenced error terms. (Our error structure *allows* for autocorrelation but does not presume it exists. If the errors were independent from period to period the first-differencing of an occupation's observations would definitely induce autocorrelation.) Instead, we utilize an approximate method to estimate $\delta_i$ directly from the observations, then subtract the estimate of $\delta_i$ from the observations for each occupation (see below).

Either autocorrelation or the absence of an occupation intercept can give rise to simultaneity bias in the reenlistment rate equation and produce inconsistent parameter estimates, and in particular, inconsistent bonus effects. Consequently, these factors must be controlled in the estimation procedure.

[illegible]

## ESTIMATION PROCEDURE

Our discussion of the estimation procedure includes the choice of functional form, weighting the observations, controlling for the occupation intercept, and controlling for autocorrelation.

### Functional Form

The functional form we estimate is the logit. For retention ($s_t$) it is a simple dichotomous logit; for reenlistment ($r_t$) and extension ($e_t$), it is a polytomous logit because the person is faced with three possible choices (reenlist, extend, or leave). With grouped data, the logit model can be approximated by making the log of the odds ratio a linear function of the explanatory variables and using generalized least squares (Berkson 1953, Weinschrott 1976). This approach has the advantage of giving ready access to inexpensive and convenient software packages that can treat such error components models as ours[*]. The logit specification also preserves the property that the reenlistment, extension, and leave rates sum to one, as required by our accounting framework.

In the dichotomous case, the dependent variable is a log transformation of the retention rate, $ln(s_{it}/(1 - s_{it}))$. Because of small cell sizes in some occupations, the retention rate may have been zero in a given period. For smoothing and to avoid taking the log of zero (Cox, 1970), $s_{it}$ is defined as $(S_{it} + .5)/(N_{it} + 1)$ where $S_{it}$ is the number of reenlisters and extenders in occupation $i$ in time period $t$ and $N_{it}$ is the total number of personnel within six months of the end of their term of service (minus those we remove as ineligible).

In the polytomous case with grouped data, the dependent variable is the log of the odds ratio relative to the base choice. We selected leaving the service as our base choice, resulting in the following dependent variables:

$ln(r_{it}/(1 - (r_{it} + e_{it}))$ for reenlistment;

$ln(e_{it}/(1 - (r_{it} + e_{it}))$ for extension.

The reenlistment and extension rates are adjusted for zero cells in the same manner as for retention. The procedure maintains the property that all three rates sum to one.

Although we do not estimate the intercepts directly, their effects can be recovered *ex post* and subsequently used in making occupation specific forecasts of reenlistment, extension, and retention rates.

Given these definitions, the coefficients of each equation represent the difference in the effect of a variable on the specific outcome versus leaving: $b_{2r} = (\beta_{2r} - \beta_{2\text{leave}})$ and $b_{2e} = (\beta_{2e} - \beta_{2\text{leave}})$. Using these coefficients one can also derive the effect of a variable on the extension rate relative to the reenlistment rate and vice versa. The effect on the extension rate relative to the reenlistment rate is simply $b_{2e} - b_{2r}$, which equals $\beta_{2e} - \beta_{2r}$. Switching the order of subtraction makes extensions the base. Also, the logit equations can be solved back to find the relationship between a variable and a rate (rather than its logit), or to find the partial derivative of a rate with respect to a variable. The partial derivatives, which we use in discussing the empirical results, have the following form:

$$\partial r / \partial X = \beta_r r(1 - r) - \beta_e re$$

$$\partial e / \partial X = \beta_e e(1 - e) - \beta_r re$$

In calculating the partial derivatives, the mean reenlistment and extension rates for the total sample provide values for r and e. The rates themselves, which are featured in the cost effectiveness discussion in Sec. VII, have the form:

$$r_{it} = \exp(\beta_r{}'X_{it} + \delta_{ri}) / D$$

$$e_{it} = \exp(\beta_e{}'X_{it} + \delta_{ei}) / D$$

where

$$D = 1 + \exp(\beta_r{}'X_{it} + \delta_{ri}) + \exp(\beta_e{}'X_{it} + \delta_{ei}).$$

## Weighting

Because the size of the eligible population varies across occupations and over time, the error terms in the reenlistment, extension, and retention equations will be heteroscedastic. Heteroscedasticity will not affect the consistency of the parameter estimates but will tend to decrease their precision. To overcome this, the observations must be weighted by weights proportional to the variance of the error terms. The estimator for the variance of the error term for the dichotomous logit model is $[1/(N_{it}s_{it}(1 - s_{it}))]$. In the polytomous case with $p_3$ as the base probability, the estimators for the variance are:

$$(p_{1it} + p_{3it}) / (N_{it}p_{1it}p_{3it}) \text{ and}$$

$$(p_{2it} + p_{3it}) / (N_{it} p_{2it} p_{3it}).$$

The weight for each observation (occupation $i$ in time period $t$) is the inverse of the square root of these estimates. We know $N_{it}$ but need estimates of the probabilities. The proportions we observe reenlisting, extending, or staying in a given occupation and period will be poor estimates because of their large variance when cell sizes are very small, as can occur in our data. Therefore, we use the proportions reenlisting, extending, and staying over all time periods ($r_i$, $e_i$, $s_i$) as our estimates of the $p_{it}$s. With small cell sizes in some occupations, this makes the weights much more stable.

## Steps in Estimation

The estimation process involved several steps. First, to eliminate the occupation intercept all variables are deviated from their *occupation-specific* means. In other words, the observations for each occupation are centered around the occupation's mean observation.[2] Second, the observations are weighted to correct for heteroscedasticity. Next, we had to obtain parameter estimates to compute residuals for the calculation of the first-order autocorrelation coefficient for each occupation, $\rho_i$. We estimated $\rho_i$ as a simple regression of the Markov process (see Eq. (2)), namely:

$$\rho_i = \Sigma e_{it} e_{it-1} / \Sigma e^2_{it-1}.$$

The parameter estimates themselves were computed by means of ordinary least squares for the retention equation and by the seemingly unrelated regressions method for the reenlistment and extension equations.[3] Fourth, to eliminate autocorrelation we used the $\rho_i$ to quasi-difference the observations. In this step, $\rho_i$ times the observation in

[2] [illegible] (1963) proves this method provides unbiased estimates of all parameters—i.e., the same estimates as if all the occupation intercepts had been included.

[3] We found in experimentation that the Seemingly Unrelated Regressions (SUR) method produced estimates about the same as those produced by single regressions for each equation, and the standard errors were basically unchanged. Generally, if the explanatory variables are identical for both equations, there is no gain in efficiency from SUR. Our equations have the same set of raw explanatory variables, however, because we weight the data with different weights for reenlistment and extension, the transformed explanatory variables are not identical between equations. The lack of increased efficiency may be due to our elimination of the intercept, for when the reenlistment and extension equations are formulated as permanent component models the error correlation between the equations is primarily due to the permanent component in the error term. Removal of that element from each equation's structure greatly reduced the correlation between the two errors. This result is important because software limitations did not permit correlation between the errors of the equations in the final step in our estimation procedure.

period $t - 1$ is subtracted from the observation in period $t$. The final step was to run a variance components model on the transformed data. This technique controls for the remaining portions of the error term, $\epsilon$ and $\xi_{it}$, and provides consistent estimates of the parameters of the reenlistment, extension, and retention logits.

It follows from these conditions that higher bonuses must increase expected manyears of service in an occupation. To begin, suppose the reenlistment rate and extension rate changes just offset each other, leaving the retention rate unchanged ($de = -dr$). The change in manyears then becomes

$$dMY = N(y_r - y_e)dr,$$

which must be positive because $y_r > y_e$ and $dr > 0$. *The increase in the occupation's expected manyears results from a shift of personnel* from extension to reenlistment. Because the retention rate remains constant, the same number of people stay in the occupation; but they commit to longer terms of service, hence to greater manyears. Manyears grow by still more when the occupation's retention rate is allowed to increase. The change in manyears partitions into the preceding *shift effect* and a *retention effect*. To see this, add and subtract $Ny_e dr$ in the equation for $dMY$:

$$dMY = N(y_r dr + y_e de) + Ny_e dr - Ny_e dr$$

$$= N(y_r - y_e)dr + N(y_e dr + y_e de)$$

$$= N(y_r - y_e)dr + Ny_e(dr + de)$$

The first part of the equation reflects the shift from extension to reenlistment, holding retention constant. The second part shows the additional increase in manyears accompanying the increase in the retention rate. Because both $dr$ and $(dr + de)$ are positive in response to a bonus increase, this demonstrates that higher bonuses increase expected manyears.

Empirically, probably more of the bonus effect on manyears comes from the shift effect than the retention effect. Table 4 shows that the effect of higher bonuses on the reenlistment rate is about three times larger than the effect on the extension rate and in the opposite direction. If we substitute for $de$ with $-.33dr$, we obtain:

$$dMY = N(y_r - y_e)dr + Ny_e(dr - .33dr)$$

$$= N(y_r - y_e)dr + N(.67y_e)dr.$$

If $y_r$ is more than 1.67 times $y_e$, then the shift effect, $N(y_r - y_e)dr$, is larger than the retention effect, $N(.67y_e)dr$. [illegible] probably

at the reenlistment decision point ($N$) times a weighted average of the reenlistment and extension rates.

$$MY = N(y_r r + y_e e), \quad y_e < y_r.$$

The weight on the reenlistment rate ($r$) is the expected number of manyears given that the person has reenlisted, which we denote $y_r$. The weight on the extension rate ($e$) is the expected number of manyears given that the person has extended (and afterward may reenlist or leave). We assume that this weight, $y_e$, is less than $y_r$ because we expect an extender, on average, to generate fewer manyears than a reenlistee.[1] Of course, some extenders will subsequently reenlist, either in their own occupation or in another. They may have extended because they were unsure of whether to continue in service, or, although sure, perhaps they wanted to migrate to another occupation and needed to extend in order to retrain. Extenders who subsequently reenlist in their current occupation probably stay in service for about as long as, or maybe a little less than, those who reenlist immediately. Extenders in the process of changing occupations might stay the same length as those remaining in the occupation and reenlisting outright. In contrast, the extenders who leave generate perhaps a single year of service in the course of the extension.

The primary effects of a change in policy are represented by a change in expected manyears. To obtain this we take the derivative of $MY$, holding $N$, $y_e$, and $y_r$ constant, which seems reasonable for the purposes of our analysis.[2]

$$dMY = N(y_r dr + y_e de).$$

We found that bonuses increased the reenlistment rate, reduced the extension rate, and increased the retention rate (Table 4). This was true for lump sum and installment bonuses, for both the first term and the second term. The responses to higher bonuses therefore can be characterized by the conditions $dr > 0$, $de < 0$, and $(dr + de) > 0$.

[1] According to figures from the Defense Manpower Data Center, [illegible] and two times.

[2] [illegible] $N$, [illegible], and [illegible] could change depending on the extent [illegible] change under consideration. For instance, [illegible] bonus that [illegible] several reasons might reduce attrition and thereby increase [illegible] persons reaching the reenlistment decision point $N$. Also, a higher [illegible] some persons to reenlist rather than extend, and these persons [illegible] terms [illegible] three years) than would persons who already [illegible] The latter persons, in fact, might be induced [illegible] the higher [illegible] These effects might or might not offset one another. [illegible] changes in bonuses and unemployment that [illegible] changes in $N$, [illegible] are negligible.

# VII. POLICY IMPLICATIONS

We utilize the results from our analysis of reenlistment, extension, and retention rates to evaluate whether lump sum bonuses are more cost effective than installment bonuses. To make the evaluation, we must first demonstrate that these rates provide clearcut information about prospective changes in expected manyears of service, for it is preferably in terms of manyears, not rates, that the cost effectiveness of reenlistment bonuses should be judged. We also must place lump sum and installment bonuses on an equal cost basis. That being done, we discuss the evaluation itself, finding support for the notion that lump sum bonuses are more cost effective. We also discuss the extent to which higher military/civilian pay or higher reenlistment bonuses could be used to counteract the downturn in retention that tends to occur during periods of low unemployment. Such steps, if taken, would dampen the swings in military retention and help stabilize the personnel force structure.

## COST EFFECTIVENESS OF LUMP SUM VERSUS INSTALLMENT BONUSES

### The Relationship Between Rates and Manyears of Service

As a general proposition, we would like to examine the total increase in manyears of service achievable under a program of lump sum bonuses versus an equal-cost program of installment bonuses. Unfortunately, our data do not measure manyears of service subsequent to reenlistment or extension, so we cannot directly quantify the change in manyears caused by a shift from installment to lump sum bonuses, holding the overall bonus budget constant. Instead, we must attempt to draw inferences about changes in manyears from our evidence on the effect of bonuses on reenlistment, extension, and retention *rates*. We now present a simple framework showing the circumstances under which certain changes in reenlistment, extension, and retention rates unambiguously imply an increase in the expected manyears of service.

We begin with a relationship between manyears (*MY*) in an occupation and the occupation's reenlistment and extension rates. This sort of relationship holds for any occupation, so we do not use occupation subscripts. The number of manyears equals the number of personnel

## Table 7

### VARIANCE OF ERROR COMPONENTS
(Error $\epsilon_t + \xi_{it}$)

| Source of Error | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| | Reenlist | Extend | Retention | Reenlist | Extend | Retention |
| Contemporaneous ($\epsilon_t$) | .002 | .006 | .003 | .006 | .009 | .005 |
| Random noise ($\xi_{it}$) | .095 | .157 | .072 | .129 | .273 | .119 |
| Total | .097 | .163 | .075 | .135 | .282 | .124 |

Table 6

SUMMARY STATISTICS ON OCCUPATION SPECIFIC AUTOCORRELATION

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| Statistic | Reenlist | Extend | Retention | Reenlist | Extend | Retention |
| Average value of $\rho_i$ | .046 | .079 | [illegible] | .061 | [illegible] | [illegible] |
| Standard deviation of $\rho_i$ | .322 | .339 | .330 | .304 | [illegible] | .329 |

coefficient is zero for each occupation, but because we pool data across occupations and the *average* value of the coefficient is zero. Although we corrected for autocorrelation, in hindsight we probably did not have to. But if one sought to estimate occupation-specific bonus effects with data for single occupations or a small set of related occupations, autocorrelation could remain a problem. Our estimates indicate that $\rho_i$ is distributed normally across occupations within each term; thus there is a reasonable probability that the autocorrelation coefficient for a given occupation will be nonzero and a potential source of simultaneity bias.

## VARIANCES OF ERROR COMPONENTS

The final estimates of our logit regression models are based upon an error specification with two components, the contemporaneous component $\epsilon_t$ and the purely random component $\xi_{it}$. These components remain after the autocorrelation component has been removed by quasi-differencing the data. Table 7 contains the estimates of the variance of $\xi_{it}$ and $\epsilon_t$, as well as an estimate of the total error variance that is the sum of those two. Evidently, contemporaneous shocks across all occupations account for a very small portion of the total error variance, with the majority being attributable to random noise.

Table 5

COEFFICIENT ON INSTALLMENT BONUS BY METHOD OF ESTIMATION
(t-statistic)

| Method of Estimation | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| | Reenlist | Extend | Retention | Reenlist | Extend | Retention |
| SUR; weighted data[a] | .0173[b] | −.0203 | .0077 | −.1085[c] | −.1144[c] | −.1082[c] |
| | (2.55) | (−1.86) | (1.28) | (−15.22) | (−8.74) | (−16.10) |
| SUR; weighted, centered data[d] | .0759[c] | −.0275[b] | .0490[c] | .0323[c] | −.0116 | .0180[b] |
| | (11.34) | (−2.64) | (7.63) | (4.00) | (−0.93) | (2.45) |
| GLS; weighted, centered, quasi-differenced data[e] | .0737[c] | −.0280[c] | .0490[c] | .0483[c] | −.0188 | .0310[c] |
| | (11.92) | (−2.61) | (9.09) | (6.85) | (−1.57) | (5.20) |

[a]Adjusts for heteroscedasticity.
[b]Significant at the 95 percent confidence level (critical t = 1.96).
[c]Significant at the 99 percent confidence level (critical t = 2.58).
[d]Also adjusts for occupation intercept.
[e]Also adjusts for autocorrelation.

term namely, an apparent shift in the pattern of second term bonus allocation during the latter periods in our data.

The pattern of change in the bonus amount coefficient as we change methodologies is the same for the retention regression as for the reenlistment regression. This is not surprising because in both instances we expected a positive, true coefficient with a downward bias because of simultaneity.[7]

The results in Table 5 suggest that the occupation intercept is a more potent source of simultaneity bias than the autocorrelation coefficient. We infer this by observing generally greater changes in the bonus amount coefficients after centering the data than after quasi-differencing the data (given that they have been centered). Theoretically, autocorrelation would not be important if the expected value of the autocorrelation coefficient were zero. This turns out to be close to what we find, as Table 6 shows, not because the autocorrelation

[7]As a side point, the full regression results for weighted, centered data (App. E) also suggest that this method improves our wage estimates. In the SUR (weighted only) results (App. D) the wage coefficients are actually negative. When the data have been centered by occupation, the wage coefficients become positive. We are not sure of the reasons for this change. Perhaps the wage coefficient is affected by its own simultaneity bias; increases in the relative military wage might be more likely when retention rates are falling overall.

heteroscedasticity, then estimates the logit regressions by SUR.[7] The procedure, which offers no control for simultaneity bias, provides a benchmark. The second procedure also weights the data and, further, centers the data for each occupation with respect to the occupation's average observation for the sample period (mid-FY76 through FY81). That is, the average observation is subtracted from each of the occupation's 11 observations. This procedure controls for heteroscedasticity and eliminates the occupation's intercept. We estimate this model also by means of the SUR technique. The third procedure does as much and controls for the occupation's autocorrelation--the possibility that the error may be correlated from period to period. First the results from the second procedure are used to estimate the autocorrelation coefficient for each occupation, then the occupations, which have already been weighted and centered, are quasi-differenced to remove the autocorrelation from the error term. We estimate this model by Generalized Least Squares (GLS), using the error component specification described in Sec. V. The estimated variances of the error components are presented in the next subsection below.

Table 5 summarizes the results of the three procedures, for brevity focusing on the coefficients for bonus installment amount. The complete logit results appear in Appendixes B, D, and E.

The first method (SUR, weighted data) indicates a positive bonus amount coefficient for the first term reenlistment logit but a *negative* coefficient for the second term. Use of the second method (SUR, weighted, centered data) changes the results: The first term coefficient becomes several times larger, and the second term coefficient reverses from negative to positive. Movement to the third method produces little additional change in the first term and a slight increase in the second term.

We expected the bonus amount coefficient in the extension logit to become more negative after the data was centered because the simultaneity problem would bias the coefficient toward zero. In the first term, the second method does make the coefficient more negative, and the third method more so. However, in the second term the change from the first to the second method makes the coefficient *less* negative, although it becomes somewhat more negative when the third method is applied. Perhaps the unexpected pattern in the second term originates from the same problem preventing us from estimating a significant effect of lump sum bonuses over installment bonuses for the second

[7]SUR allows for the possibility of correlation between the reenlistment and extension logit regressions, and, as noted in Appendix C, almost no such correlation was found.

## Percent Nonhigh School and Percent Black

Although the percent nonhigh school is statistically significant for the first term extension rate and all three second term rates, the derivatives for this variable are small. Even reasonably large swings in the percent nonhigh school therefore have little effect on the various rates. The practical insignificance of this variable may stem from its referring to education level at the reenlistment point, not at the initial service entry point. Many personnel who enter the service without a high school diploma manage to obtain a GED by the end of their first term. In our sample they would be counted as having a high school diploma or GED, so the sample blurs any behavioral distinction between persons who complete formal high school and those who do not. Such a distinction appears in the analysis of first term attrition. For example, Buddin (1984) finds GED attrition patterns that follow those of nonhigh school graduates rather than those of diplomates. Chow and Polich (1980) find high school graduation to be a significant determinant of an individual's first term reenlistment behavior; the reenlistment probability is lower for high school diplomates.

Across the board, blacks appear more likely to remain in service. We estimate nearly identical effects for the rates in the first and second term. Moreover, a 10 percent change in the percent black can have a noticeable effect on retention behavior; the reenlistment rate would rise by 1.1 percentage points, the extension rate by .3 percentage point, and the retention rate by about 1.4 percentage points. In contrast, a 10 percent increase in the percent nonhigh school would produce changes in the first term rate of .2, −.2, and .1 percentage points respectively.

## Simultaneity Bias

The econometric section proposes a two equation model of rate determination and bonus determination. The model identifies two potential sources of simultaneity bias in the rate equation, and by implication, failure to correct for the bias would lead to downward-biased bonus coefficients. The sources are the occupation's intercept $\delta_j$ and its autocorrelation coefficient $\rho_j$. The results presented above in the form of derivatives come from our final model, which attempts to control both sources of simultaneity bias. However, to grasp the importance of controlling for each of the sources, we now present bonus amount coefficients derived from three different estimation procedures. The first procedure simply weights the data to correct for

postponing departure from service, then there would not be an increase in expected manyears of service, even though an increase in military pay would increase today's retention rate. Instead, the number of personnel with a temporary attachment to service would increase, which might complicate service personnel managers' tasks in planning for rotations and promotions. Higher bonuses might relieve this situation by inducing reenlistments, not extensions.

## Unemployment Rate

We hypothesized that higher unemployment would increase the reenlistment and retention rates, and our results confirm this. The results also show a negative effect of unemployment on the extension rate. The magnitude of the unemployment effect on the reenlistment rate is several times the effect on the extension rate.[5] In addition, the pattern of response to a higher unemployment rate parallels that of a higher reenlistment bonus. By implication, the consequences of a *decline* in the unemployment rate, namely a lower reenlistment rate and a higher extension rate, could be offset to some extent by higher bonuses and, implicitly, by more extensive bonus coverage across military occupations.

On net, higher unemployment leads to a higher retention rate. Moreover, second termers appear to be about three times more responsive than first termers. An increase in unemployment from 7 percent, say to 8 percent, causes a three percentage point increase in the second term retention rate versus a one percentage point increase in the first term rate. This may indicate greater caution at the second term reenlistment point in evaluating the benefits and costs of leaving the military to seek a civilian career. Compared with first termers, second term personnel probably are more skilled and have skills of a more specialized nature, thus an acceptable civilian job may be more difficult to locate. Marriage and family obligations probably also play a larger role among second term personnel. In any case, our findings suggest that, relative to first termers, second termers tend to time their departure from service to coincide with good national employment conditions.

[5] Goldberg and Warner (1982) found similar results for first term Navy skill groupings. Unemployment effects on extensions were negative in 7 of their 9 skill groups and smaller than the positive effects on reenlistments.

sample, the percent of second term occupations covered by bonuses rose from 22 percent at the end of FY79 to 28 percent in FY80 and then to 42 percent in FY81. The econometric procedure does not control for the added simultaneity resulting from such a change in bonus allocation policy. Fortunately, there appears to be no similar change in first term bonus allocation policy.

## Military/Civilian Wage Index

We hypothesized that a higher military/civilian wage index would increase the reenlistment, extension, and retention rates. The results confirm these hypotheses for both the first and second term, with the exception of the second term reenlistment wage effect, which is positive but significant only at about the 85 percent confidence level. Also, as we observed for bonuses, the derivatives for the two terms are similar. For the reenlistment rate, the wage derivatives for the first and second term are .0029 and .0022, the extension rate derivatives are .0051 and .0055, and the retention rate derivatives are .0080 and .0079. Thus a 1 point increase in the military/civilian wage index will increase retention in either term by .8 percentage points (a 3.8 percent increase in retention for the first term and a 1.7 percent increase for second term). As expected, the retention rate derivatives equal approximately the *sum of the reenlistment and extension rate derivatives.*

In formulating the hypotheses, we did not expect that the effect of the wage on the extension rate would exceed its effect on the reenlistment rate. However, the results show an extension effect about *double* in size to the reenlistment effect, which implies that the extension category attracts personnel whose decision to remain in service is more strongly influenced by pay considerations. This raises a question requiring further research: Does the proportion of extenders who subsequently choose to reenlist, rather than to leave, rise or fall as relative military pay increases? Perhaps today's increase in military pay convinces more people to remain in service and to seek a military occupation they perceive to be more satisfying than their current one. If so, a natural avenue would be to extend to retrain, and then to reenlist. Alternatively, higher military pay may encourage delayed exit; rather than leave immediately, some personnel may stay a little longer, extending and afterward leaving.

Of course, these alternative explanations carry different consequences for the number of *manyears* generated by a military pay increase. The greater the extent of retraining and reenlisting among extenders, the higher the expected manyears of service. If the extent of retraining were low and extensions were mainly a vehicle for

The results for the first term confirm the hypothesis [illegible] lump sum bonuses exert a greater effect on the reenlistment rate [illegible] installment bonuses. A $1000 increase in a lump sum bonus [illegible] approximately equal to a 1/2 step increase in the [illegible] will increase the reenlistment rate by 1.25 percentage points [illegible] similar increase in an installment bonus will produce [illegible] percentage point increase.[4] Although these changes may seem small, the lump sum effect represents a 9 percent increase in the average reenlistment rate of our sample, but the installment effect represents only a 6 percent increase. Lump sum bonuses are almost twice as effective in reducing the extension rate: A $1000 increase in a lump sum bonus reduces the extension rate by .46 percentage points, but a similar installment bonus increase produces only a .26 percentage point fall. The table shows a modest increase (.98 percentage points for a $1000 increase in lump sum and .81 for installment) in the retention rate attributable to lump sum bonuses. The t-statistics in row 3 of Table 4, indicate that the lump sum bonus effects are significantly different from the installment effects. The reenlistment and extension effects are significant with 99 percent confidence, and the retention effect is at 90 percent.

The second term results reveal no differential effect of a lump sum bonus on the reenlistment rate. However, we observe a large negative effect on the extension rate about equal in size to that of the first term. The zero reenlistment effect and the negative extension effect together produce a negative effect on retention, which contradicts our hypothesis. In other words, the effect of second term reenlistment bonuses was lower after the shift to lump sum payment occurred, .0042 rather than .0077 before April 1979.

The unexpected results for the second term probably stem from an inadequate adjustment for simultaneity bias during the lump sum periods in our data. The rapid expansion in second term bonus use from mid-FY79 through FY81 apparently signals a structural shift in the bonus allocation procedure. Occupations with low retention rates that had not previously received bonuses began to receive them. In our

[4]These results are similar to Goldberg and Warner's (1982) findings for Navy skills. They used a polytomous logit model to estimate reenlistment and extension equations for [illegible] separate skill groupings. Their bonus variable was included as part of the stream of future military pay and therefore constrained to have the same effect as [illegible] in other military income. Despite this difference in specification, their estimates suggest that a one step increase in an installment bonus multiple would increase the first term reenlistment rate by 1.3 to 2.3 percentage points. Our estimates [illegible] DoD as a whole, a one step increase would bring a 1.8 percentage point increase in the first term reenlistment rate. For lump sum bonuses, Goldberg and Warner find [illegible] step increase in the bonus multiple would increase Navy first term reenlistment [illegible] percentage points. Our model, by comparison, would predict [illegible] percentage point increase in DoD wide first term reenlistment rates.

derivative, negative extension rate derivative, and positive retention rate derivative, all statistically significant for each term, support the hypothesis for each bonus type.[2]

## Lump Sum Versus Installment Bonus Effects

The regression specification tests directly whether the effects of lump sum bonuses dominate those of installment bonuses, as hypothesized. The entries in the third row of Table 4 (Lump sum-Installment effect) are the *difference* between the derivatives for lump sum and installment bonuses. A positive entry means that the lump sum derivative exceeds the installment derivative, and a negative entry means the reverse. For correct interpretation, one should recognize that the derivatives contrast the effects of installment and lump sum bonuses of equal face value—i.e., a $1000 installment bonus change versus a $1000 lump sum bonus change. Because installment bonuses are paid over time, the present value of an installment bonus will be less than its nominal value, whereas the present value of a lump sum bonus equals its nominal value. Given this difference in value and the greater riskiness of installment bonuses (Sec. III), one expects the results for lump sum bonuses to dominate those of installment bonuses.[3] The policy question, addressed in the next section, is whether *lump sum bonuses perform sufficiently better* to be judged more cost effective than installment bonuses.

Because the drafting of young men into the military did not end until December 1972, first term individuals with four to five years of service in the first period of our data (March to September 1976) entered the military during the draft era. Some of those were undoubtedly draft-induced enlistments, and their retention behavior patterns may be different from those of volunteers. Presumably such individuals would be less responsive to bonuses as an effort to increase retention. To see whether inclusion of these individuals may have biased our bonus results, we reran the model excluding the first period of observation for each skill. The bonus coefficients (installment and lump sum) did not change much in the *reenlistment* equation and were slightly larger in absolute value in the extension equation. The overall effect on retention was no change in the installment coefficient and a slight decrease in the lump sum coefficient. Acknowledging the general stability in the bonus coefficients, it appears that the inclusion of some draft-induced enlistments in our earliest period of observation does not affect our results.

An alternative approach would be to discount the installment bonus to place it in present value terms. Such discounting appears to make the installment bonus directly comparable to the lump sum bonus. However, the appearance could be deceiving: It is not clear *a priori* what discount rate to use, yet the choice of the discount rate affects the size of the coefficient on the discounted installment bonus. Further, lump sum and installment bonuses may not be equally risky assets; expressing them both in present values would not remove any intrinsic differences in riskiness. Our approach allows perceived differences in present value and riskiness to be reflected in the estimated coefficients, not imposed ahead of time.

## Table 4

FIRST DERIVATIVES
(Approximate t-statistic)[a]

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| Variable | Reenlist | Extend | Retention[b] | Reenlist | Extend | Retention |
| Installment bonus | .0091[c] | −.0026[c] | .0081[c] | .0117[c] | −.0034[c] | .0077[c] |
| ($1000) | (12.81) | (−3.82) | (8.62) | (7.41) | (−3.25) | (4.93) |
| Lump sum bonus | .0125[c] | −.0046[c] | .0098[c] | .0111[c] | −.0058[c] | .0042[c] |
| ($1000) | (17.50) | (−7.51) | (10.2) | (9.91) | (−8.13) | (3.36) |
| Lump sum − | .0034[c] | −.0020[c] | .0017 | −.0006 | −.0024[c] | −.0035[c] |
| installment effect[d] | (4.79) | (−2.82) | (1.68) | (−0.46) | (−2.70) | (−2.62) |
| Military/civilian | .0029[c] | .0051[c] | .0080[c] | .0022 | .0055[c] | .0079[c] |
| wage index | (4.69) | (11.64) | (10.54) | (1.50) | (6.89) | (5.39) |
| Unemployment rate | .0140[c] | −.0040[c] | .0104[c] | .0322[c] | −.0051[e] | .0298[c] |
| in previous period | (9.03) | (−3.69) | (5.47) | (7.89) | (−2.28) | (7.33) |
| Percent without HS | .0002 | −.0002[c] | −.0001 | −.0014[c] | .0007[c] | −.0006 |
| diploma or GED | (1.85) | (−4.20) | (−1.31) | (−4.48) | (4.41) | (−1.88) |
| Percent black | .0011[c] | .0003[c] | .0014[c] | .0017[c] | .0003[c] | .0015[c] |
| | (13.74) | (5.00) | (13.56) | (7.26) | (3.63) | (10.28) |

[a]See Appendix B for the regression results and Appendix C for a discussion of the standard errors underlying the t-statistics.

[b]Entries are based on retention equation estimates from the dichotomous logit model.

[c]Significant at the 99 percent confidence level (critical t = 2.58).

[d]The entries on this row should be read as the difference between the first derivative with respect to a lump sum bonus and the first derivative with respect to an installment bonus: $\partial P/\partial B_{\text{Lump Sum}} - \partial P/\partial B_{\text{Installment}}$.

[e]Significant at the 95 percent confidence level (critical t = 1.96).

derivatives, and Appendix C describes the method for computing the derivatives' t-statistics.

### Bonus Amount

The bonus amount effects are represented by the derivatives with respect to the installment and lump sum bonus variables. By hypothesis, a higher reenlistment bonus of either kind should induce more personnel to remain in service in their bonus-covered occupation and to sign on for longer obligations. The positive reenlistment rate

# VI. EMPIRICAL RESULTS

On the whole, the results of our empirical analysis strongly confirm our hypotheses. Higher bonuses do increase the reenlistment rate, decrease the extension rate, and increase the retention rate. The same pattern is found for the unemployment rate, and the military civilian wage index has a positive effect on all three rates, again as expected. At the first term reenlistment point, lump sum bonuses appear to be more potent than installment bonuses of equal nominal size, thus confirming a key hypothesis of this study. However, at the second term reenlistment point, the lump sum bonus effects are mixed.

The section presents the empirical results from a different approach than generally used in discussions of logit regression results. A *polytomous* logit model raises difficult problems in interpreting the regression coefficients as tests of hypotheses. To circumvent these problems, we convert the regression results to first derivatives that explicitly show the change in a rate with respect to the change in an explanatory variable. Appendix I discusses the use of first derivatives in testing hypotheses given a logit model.

The section also presents results with respect to simultaneity bias in the bonus coefficient. The bias is substantial but our tactics for controlling it seem to be effective. The section concludes with estimates of the variances of the error components, which were described in the econometric model section. We find little evidence for the importance of contemporaneous shocks affecting all occupations at any one time.

## PARTIAL DERIVATIVES AND TESTS OF HYPOTHESES

Table 4 displays the partial derivatives and their t statistics. An overview indicates that *nearly all the hypotheses have been confirmed for both the first term and the second term.* The size and statistical significance of the derivatives depends on the point of evaluation, and we have chosen to evaluate the derivatives at the mean reenlistment, extension, and retention rates observed over the periods covered by our data, mid-FY76 through FY81. Evaluation at other levels observed during our data period would produce largely the same results. Appendix B contains the logit regression coefficients underlying the

The mean rates used for first term were .139 for reenlistment, .076 for extension, and [illegible] for retention; second term rates were .354, .101, and [illegible], respectively.

exceeds $s_r$ by a factor of two, the shift factor dominates the increase in retention because of an increase in bonus.[3]

The expression for $dMY$ shows that the size of the increase in manyears depends on the size of the reenlistment rate increase ($dr$) and the size of the retention rate increase ($dr + de$). The empirical results for the first term show lump sum bonuses producing greater increases in these rates than installment bonuses of equal nominal size. *By implication, first term lump sum bonuses produce greater expected manyears in an occupation than installment bonuses.* Although this is a necessary condition for lump sum bonuses to be more cost effective, it is not a sufficient one.[4] The empirical results for the second term, by comparison, show no significant difference in the reenlistment rate and a lower retention rate for lump sum bonuses. At the second term, then, lump sum bonuses appear to produce fewer expected manyears, implying that second term lump sum bonuses cannot be more cost effective than installment bonuses. However, we are not confident in the second term lump sum results because of the major expansion in second term bonus coverage occurring in the periods coincident with the payment of bonuses by lump sum (Sec. VI). For this reason the remainder of the present section concentrates on the first term results.

## Placing Bonuses on an Equal Cost Basis

The empirical results quantify the differential effect of lump sum versus installment bonuses on reenlistment, extension, and retention rates. The bonus variables were expressed in nominal amounts (e.g., a $4000 lump sum bonus or a $4000 installment bonus), and although that seems natural from the service member's viewpoint, we now must shift to the viewpoint of the services bearing the cost of the bonuses.

[3] We continue to examine expected manyears in an occupation [illegible] an occupation's manyears do not necessarily imply equivalent [illegible] servicewide manyears because of migration across occupations. At one extreme [illegible] extenders [illegible] occupational migration [illegible] suppose there were [illegible] Then decreasing an occupation's ext[illegible] result in greater expected manyears in the occupation but not in the service [illegible] not [illegible] the personnel would have migrated and supplied their [illegible] discussed at various points in the text, our extension [illegible] well as personnel [illegible] about [illegible] attract both kinds of extenders, red[illegible] overall service manyears, but reenlist[illegible] higher bonuses [illegible] greater expect[illegible]

[4] [illegible] is more cost [illegible] installment [illegible] cost effectiveness [illegible] whether the additional [illegible] generated [illegible] bonuses are sufficiently [illegible] than their additional [illegible]

From this perspective, a \$4000 installment bonus clearly costs less than a \$4000 lump sum bonus. Only a part of the installment bonus is paid each year, and the unpaid portion can in effect be invested, with the interest earned being applied to meeting the future installment payments. Further, not all personnel who reenlist finish their term, so some installment payments will not have to be made. Consequently, a given bonus budget, if expended in the form of installment bonuses, can fund either greater bonus coverage across occupations or higher bonuses, or both. For simplicity we assume that bonus coverage is held constant.[5] With bonus coverage constant, the bonus budget can be funneled entirely into higher installment bonuses.

To determine how much higher, we denote $B_L$ as the amount of lump sum bonus, $B_I$ as the amount of installment bonus, $r$ as the government interest rate, and $p$ as the probability of continuing from one year of service to the next. Reenlistees are assumed to select a four year term of service.[6] Now the cost of a lump sum bonus for a 4 year term equals the amount of the bonus itself, $B_L$. This amount could instead be used to fund an installment bonus of amount $B_I$. The installment bonus would be paid in four equal annual installments, beginning at the start of the reenlistment term. The amount of each installment payment would thus be $.25B_I$. The size of the installment bonus that could be funded by the lump sum bonus may be determined as the solution to:

$$B_L = .25\,B_I(1 + p/(1 + r) + p^2/(1 + r)^2 + p^3/(1 + r)^3)$$

Observe that because of discounting and attrition, the cost equivalent installment bonus $B_I$ exceeds the lump sum bonus $B_L$. (The formula changes somewhat when the possibility of recouping the "unearned" portion of lump sum bonuses is taken into account; see App. F.)

Based on the formula, Table 8 shows the size of a four year installment bonus that could be provided at the same cost as a \$1 lump sum bonus. The values in the table vary from about 1.10 to 1.20 in relation

[5] This assumption will have little effect on the cost effectiveness comparison because the increase in expected manyears from being able to pay higher installment bonuses, bonus coverage constant, is about the same as the increase in expected manyears from increasing coverage but keeping installment and lump sum bonuses equal in nominal size.

[6] We could instead have chosen a three year term of service, or a mix of three, four, five, and six year terms. The results of the comparison are unaffected by this choice so long as the distribution of terms chosen does not differ between the two kinds of bonuses.

## Table 8

COST EQUIVALENT INSTALLMENT BONUS
PER $1 LUMP SUM BONUS
(Assumes a four year term of service)

| | Continuation Rate | |
|---|---|---|
| Interest Rate | .95 | .97 |
| .04 | 1.14 | 1.11 |
| .06 | 1.17 | 1.14 |
| .08 | 1.20 | 1.17 |
| .10 | 1.23 | 1.20 |

to the assumed values of the government interest rate and the continuation rate.[7] In keeping with our having adjusted the bonus amounts in the empirical work for inflation, the interest rate should be interpreted as the "real" rate—that is, the inflation-adjusted rate. (The nonadjusted rate would be higher by the amount of the annual rate of inflation.) The values of the continuation rates accord with actual experience, the lower rate (.95) reflecting continuation experience in the second term, and the higher one (.97) reflecting the third term.[8]

The usefulness of Table 8 lies in converting lump sum bonuses into installment bonuses of equal cost. For example, a lump sum bonus of $3700 would correspond to an installment bonus of $4440 (1.20 × $3700), given an interest rate of .08 and a continuation rate of .95. The figure of $3700 typifies the level of first term reenlistment bonuses paid during our period of analysis, FY76–FY81, expressed in FY76 dollars (see the average bonus amount in App. A). Of course, actual bonus amounts vary from occupation to occupation depending on the

[7]If recoupment of lump sum bonuses from early leavers were effective, then the values in the table would range from about 1.05 to 1.15, as seen in App. F. This would decrease the cost advantage of installment bonuses, but recoupment might also reduce the behavioral response to lump sum bonuses. Apparently little recoupment of reenlistment bonuses occurred during our sample period. As suggested in Sec. III, the low probability of recoupment may have made lump sum bonuses seem less risky than installment bonuses, and so may have contributed to the greater responsiveness we estimated for lump sum bonuses. Greater recoupment of lump sum bonuses might decrease this responsiveness and so perhaps not lead to still greater cost effectiveness. The effect on responsiveness would probably depend on the recoupment policy; there might be little effect if recoupment were mandatory only in cases of disciplinary actions, and not in cases of injury resulting in physical disability, for example.

[8]The continuation rates are representative values selected from DoD level tabulations for years FY76 through FY82 for personnel between reenlistment decision points.

bonus step and from person to person depending on the pay grade and the length of reenlistment.

### Cost Effectiveness Comparisons

The foregoing provides a framework for relating rates to expected manyears and for converting lump sum bonuses to cost equivalent installment bonuses. Suppose a $1000 increase in lump sum bonus $B_L$ increased the first term reenlistment rate by $dr_L$ and the first-term retention rate by $(dr_L + de_L)$. The same amount of money, if disbursed as an installment bonus, would create increases of $dr_I$ and $(dr_I + de_I)$. For compactness we write the changes in the retention rate as $ds_L$ and $ds_I$ respectively. Again, the changes in the reenlistment and retention rates correspond to *cost equivalent* changes in the lump sum and installment bonuses.

The change in manyears associated with these changes can be expressed as:

$$dMY_L = (y_r - y_e)dr_L + y_e(ds_L)$$

and

$$dMY_I = (y_r - y_e)dr_I + y_e(ds_I).$$

The lump sum bonus will yield a larger increase in expected manyears in an occupation if the difference in the two expressions is positive. The difference

$$dMY_L - dMY_I = (y_r - y_e)(dr_L - dr_I) + y_e(ds_L - ds_I)$$

will be positive if the reenlistment and retention rate increases caused by the lump sum bonus exceed those of the installment bonus. Whether this condition is met depends on $(dr_L - dr_I) > 0$ and $(ds_L - ds_I) > 0$. The partial derivatives for bonuses (Table 4) can be used to infer the various rate changes that would result from cost equivalent bonus changes. These changes appear in Table 9. In constructing Table 9 we assumed that a $1000 lump sum bonus would translate into a $1200 installment bonus, an amount lying at the upper end of the values in Table 8.

*Table 9 confirms that lump sum bonuses are more cost effective than installment bonuses, at least at the first term.* For cost equivalent increases, the lump sum effect exceeds the installment effect on reenlistment and equals the installment effect on retention. When viewed in relation to the expression for net change in manyears, the results also imply that the added advantage of the lump sum bonus comes

Table 9

PARTIAL DERIVATIVES FOR COST EQUIVALENT CHANGES IN LUMP SUM AND INSTALLMENT BONUSES

| Change | Change in Reenlistment Rate | Change in Extension Rate | Change in Retention Rate |
|---|---|---|---|
| \$1000 increase in lump sum bonus | .0125 | −.0046 | .0098 |
| \$1200 increase in installment bonus | .0109 | −.0031 | .0097 |
| Difference in change: (Lump sum − installment) | .0016 | −.0015 | .0001 |

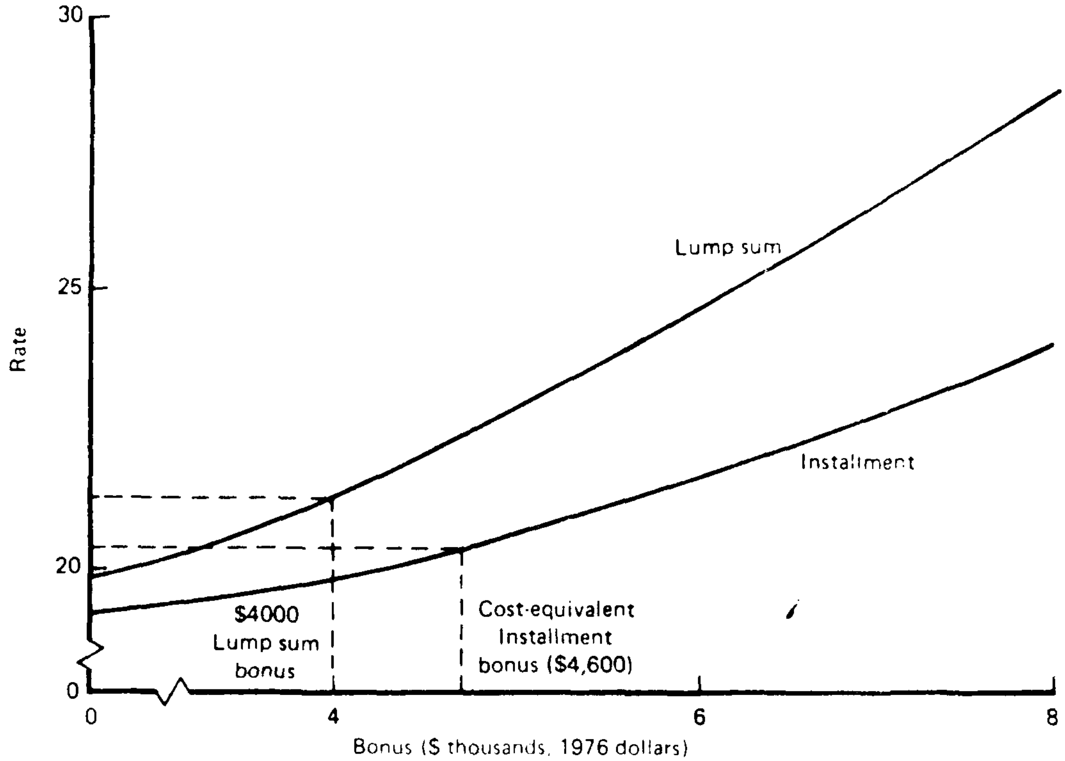
mainly from its ability to shift personnel from extension to reenlistment, rather than to retain a larger number of personnel.

To facilitate bonus comparisons over broader or more diverse changes in bonus amounts, Figs. 1 and 2 depict the predicted values of *first term reenlistment, extension, and retention rates for lump sum* and installment bonuses. (The predictions are based on our regression analysis.) Other explanatory variables—the military/civilian wage index, the unemployment rate, and the percents black and nonhigh school—have been held at their means. Also, the predictions have been centered to correspond to the average rates, rather than the rates in any particular occupation, during our period of analysis.[9]

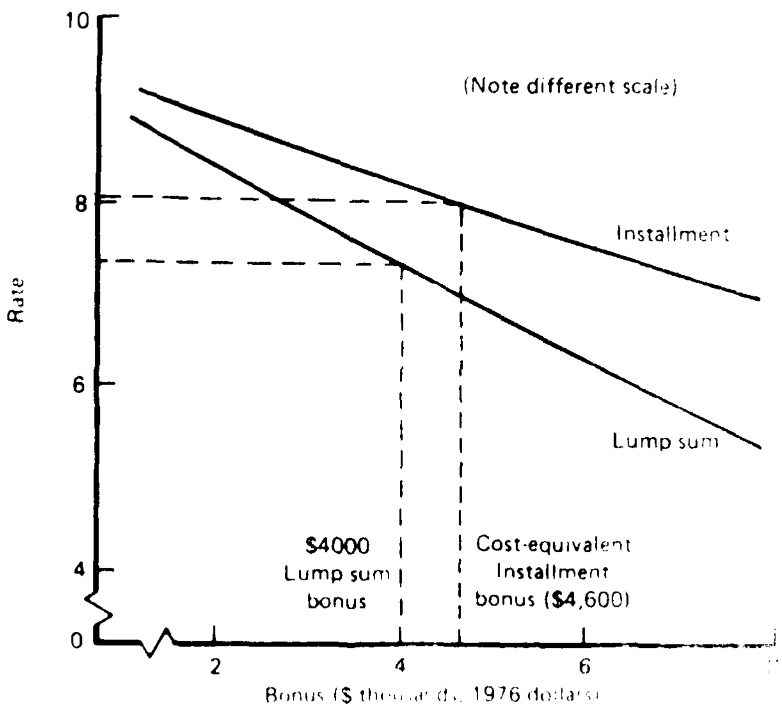
The figures may be used to estimate the rates attainable with a given lump sum bonus and its cost equivalent installment counterpart. For instance, a \$4000 lump sum bonus (slightly above the average bonus in FY76 dollars) would be equivalent to an installment bonus of \$4400-\$4800, depending on the interest rate and continuation chosen (Table 8).[10] Thus, reenlistment, extension, and retention rates can be read from the figure for the \$4000 lump sum bonus and, say, a \$4600 installment bonus; we have drawn in the lines as a guide. The figures

[9]The use of our empirical results for the purpose of predicting and forecasting will be addressed in a subsequent report.

[10]A \$4000 bonus is about equal to the average bonus paid during our analysis period, in FY76 dollars. To adapt the table or figures to other fiscal years, bonus amounts can be rescaled from FY76 dollars to the year chosen by means of, say, the Consumer Price Index. That index rose approximately 85 percent from 1976 to 1984.




(a) Reenlistment rate




(b) Extension rate

Fig 1 Predicted first term reenlistment and exten[illegible] rates for lump sum and installment bon[illegible]

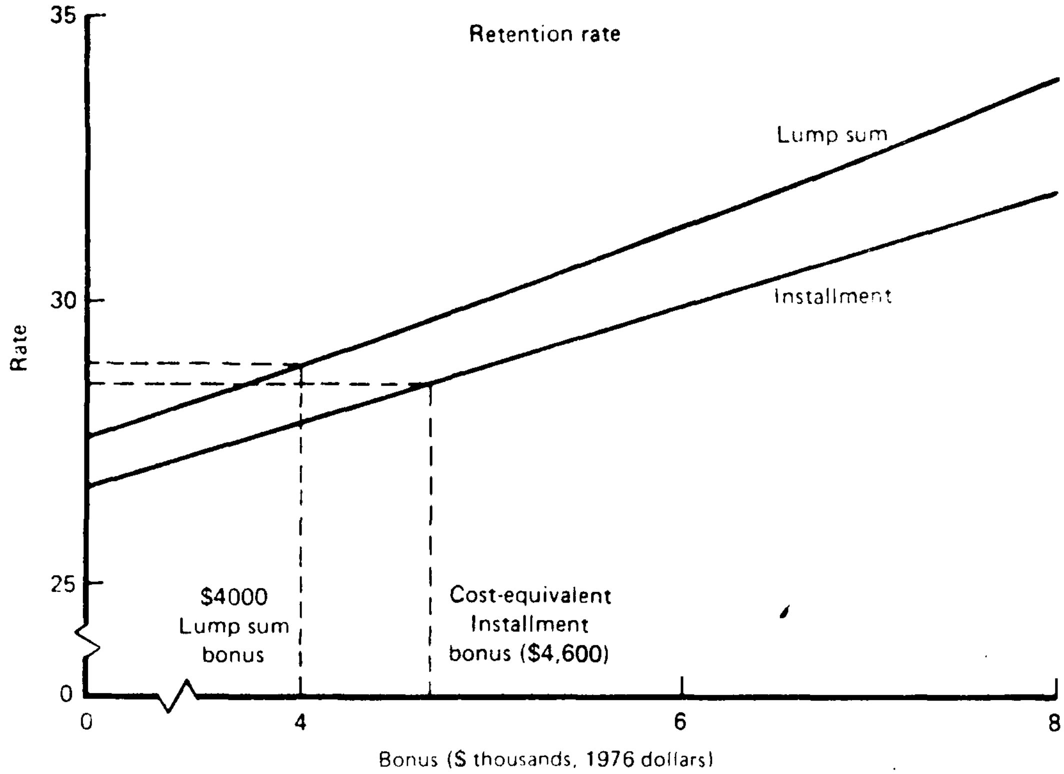



Fig. 2—Predicted first term retention rate
for lump sum and installment bonuses

permit an immediate comparison of the strength of the bonus effects, and these can in turn be judged in the context of our manyears equation. *The figures not only verify that lump sum bonuses are more cost effective than installment bonuses, but show that the dominance of lump sum bonuses becomes greater at higher bonus amounts.*

In addition, Figs. 1 and 2 illustrate the extent to which higher bonuses can alter the three rates, and these relationships are interesting by themselves. With this in mind, we now turn to a comparison of how well bonuses and wage changes fare in counteracting the effects of lower unemployment on retention behavior.

## MITIGATING THE EFFECTS OF LOWER UNEMPLOYMENT

The nation's recovery from the recession of FY82-83 brought sharp declines in the unemployment rate. Unemployment was over 10 per-

cent in mid-FY83, but it had fallen to under 8 percent by mid-FY84. As seen in Table 4, the effect of a change in the unemployment rate is quite large, especially for the second term. A two point drop in the unemployment rate results in a 2 point drop in the first term retention rate and a 6 point drop in the second term retention rate, other things equal. The fall in retention is due to the fall in the reenlistment rate; were it not for the slight rise in the extension rate, the fall would have been even greater. Thus, as the economic conditions improve, retention rates will decline. What can the services do to counteract that fall?

Higher military wages and bonuses increase retention and therefore appear to be likely policy tools to offset the detrimental effects of lower unemployment rates. In this section, we compute the size of the increase in the military/civilian wage index or reenlistment bonuses needed to counteract lower unemployment and discuss the implications for expected manyears of service. We do not, however, consider the issue of the cost effectiveness of a change in military pay relative to a change in bonus amount in improving retention rates. That issue is a much more complex one than we have evidence to address.

The following examples pertain to skills that offer a bonus. A pay increase affects all skills, and the retention effects are therefore more widespread than from an increase in bonus amount in bonus skills. For comparison purposes, we assume that bonus coverage and amounts are not affected by the pay increase.[11] The initial retention rates are 27 percent for first term (decomposed as 17 percent reenlistment, 10 percent extension) and 45 percent for second term (35 percent reenlistment, 10 percent extension).[12] Suppose the unemployment rate falls by 2 percentage points from 10 to 8 percent, dropping first and second term retention rates to 25 percent and 40 percent. The objective of the example is to restore the retention rates to their original levels before the decline in unemployment by adjusting the military/civilian pay index or the bonus level. After that adjustment, we compare the composition of retention with its initial composition before the decline in unemployment.

[11] An increase in military pay actually increases the amount of [illegible] bonus because bonus amount is the product of the bonus step and monthly [illegible]. We ignore this secondary effect.

[12] Recall that we employ a six month accounting period, as discussed in Sec. II. The retention rates would be higher were a the usual twelve month accounting period used.

## Increases in the Military/Civilian Wage Index

If the military/civilian wage index is increased by 2.5 points in the first term and by 7 points in the second term, the retention rates return to their original levels of 27 percent and 45 percent. The large required pay increase for the second term results from the greater responsiveness of the reenlistment rate to unemployment, yet similar first and second term pay effects. Table 10 displays the effect of these changes on the reenlistment and extension rates. For both terms, the reenlistment rate is lower and the extension rate higher than before the decline in unemployment. The percentage of those first termers remaining on active duty who are extenders rises from 37 percent to 44 percent; for the second term, the proportion of extenders rises from 22 percent to 33 percent. Thus higher pay can offset a lower unemployment rate, but the average length of commitment—and therefore expected manyears—of those choosing to remain in the service will decline.

Table 10

ADJUSTING PAY TO OFFSET LOWER UNEMPLOYMENT
(Percent)

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| | Unemployment | | Pay increase | Unemployment | | Pay increase |
| | 10% | 8% | 2.5 points | 10% | 8% | 7 points |
| Reenlistment | 17 | 14 | 15 | 35 | 29 | 30 |
| Extension | 10 | 11 | 12 | 10 | 11 | 15 |
| Retention | 27 | 25 | 27 | 45 | 40 | 45 |

## Increases in the Bonus Level

The picture differs for reenlistment bonuses. The results, shown in Table 11, rely on a bonus effect adjusted to reflect the current hybrid payment system. Under that system, half of the reenlistment bonus is paid in a lump sum and the other half is paid in two installments. Among first term bonus skills, a $2230 increase in the bonus (in 1976 dollars) negates the effects of the decline in unemployment. The increase translates to about $[illegible] in 1984 dollars. The second term is

Table 11

ADJUSTING BONUS LEVELS TO OFFSET LOWER UNEMPLOYMENT
(Percent)

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| | Unemployment | | Increase Bonus | Unemployment | | Increase Bonus |
| | 10% | 8% | $2230[a] | 10% | 8% | $8400[b] |
| Reenlistment | 17 | 14 | 17 | 35 | 29 | 38 |
| Extension | 10 | 11 | 10 | 10 | 11 | 7 |
| Retention | 27 | 25 | 27 | 45 | 40 | 45 |

[a]About $4125 1984 dollars.
[b]About $15540 1984 dollars.

contrast, requires a much larger increase—about $8400 in 1976 dollars ($15540 in 1984 dollars)—again because of the large unemployment effect on reenlistment.[13] In the first term, the relative proportion of extenders has not changed: reenlistment and extension have returned to their previous levels. But in the second term, the reenlistment rate is several points higher and the extension rate several points lower than initially. Thus, in addition to retaining those who would have left, the huge increase in second term bonuses results in a prominent shift from extension to reenlistment.

## Comparison of Manyear Effects

Our examples illustrate the pay or bonus increase required to return the retention rate to its original level. But what about the change in manyears caused by the fall in unemployment: Does the number of expected manyears also return to the same level as before the drop in unemployment or is it different?

Let $MY_0$ represent expected manyears initially, $MY_1$ expected manyears after the unemployment decline, and $MY_2$ expected manyears after the military pay increase to offset the unemployment decline. The pay increase makes $MY_2$ greater than $MY_1$, but it

[13]The prevailing $16000 ceiling ($20000 for nuclear and special Navy skills) on bonuses would make this increase infeasible except for occupations with low initial bonus steps. In FY81, the last year of our data, the average second term bonus step ranged from 1.1 for the Air Force and 1.5 for the Army, to 2.9 for the Marine Corps and 3.4 for the Navy. FY81 first term averages were: Air Force, 1.8, Army, 1.8, Marine Corps, 3.4, and Navy 2.9. The maximum step is 6.

appears that $MY_2$ will be less than $MY_0$. Given the formula for number of manyears, $MY = N(y_r r + y_e e)$, it follows that $MY_0 > MY_2$ because $r_0 > r_2$ and $e_0 < e_2$, assuming $y_r$ and $y_e$ are constant. This result, true for both first and second term, is caused by the greater effect of pay on the extension rate than on the reenlistment rate. Thus higher pay can offset a lower unemployment rate, but the average length of commitment and the number of expected manyears will be less.

As shown earlier, an increase in the bonus amount also increases the number of expected manyears. Furthermore, the bonus increase needed to offset the decline in unemployment does *not* reduce expected manyears. Table 11 shows that when the retention rate is restored to its initial level, the first term reenlistment and extension rates also return to their original levels. By implication, expected manyears after the bonus increase equals the initial level $(MY_2 = MY_0)$. Among second term bonus occupations expected manyears actually increases. This follows from the higher reenlistment rate and the lower extension rate, given that the retention rate again stands at 45 percent.

## Implications

Bonuses appear to be an effective tool for counteracting lower unemployment in the civilian economy and do not result in a higher proportion of short commitments. Higher pay is effective, too, but an increase in pay must be applied to all skills whether they have retention problems or not. The targetability of bonuses seems preferable when retention rates need to be improved selectively. Moreover, bonuses can be temporary, but pay increases are less so. Were unemployment to rise again in our example, the bonus increases could be rescinded. Indeed, the issue may not be so much whether bonuses are flexible and effective, but whether the bonus budget provides sufficient latitude to take advantage of a situation.

The examples also demonstrate the important difference between a change in the retention rate and a change in expected manyears. A given retention rate may correspond to various numbers of expected manyears depending on the combination of reenlistment and extension rates. An increase in retention due to a change in pay may produce a different number of expected manyears than a similar increase due to a change in bonuses.

# VIII. CLOSING THOUGHTS

## MAJOR FINDINGS

By analyzing reenlistment and extension rates separately, we have obtained new information on the effects of bonuses, pay, and unemployment on retention levels and on expected manyears of service. The analysis also helps clarify the issue of simultaneity bias in the estimation of bonus effects. Our major findings with respect to these areas of analysis and their policy aspects include the following:

- At the first term reenlistment point, lump sum bonuses are more cost effective than installment bonuses in increasing expected manyears in an occupation. Further, lump sum bonuses remain more cost effective even when bonuses are recouped from those who do not complete their reenlistment term. The added advantage of lump sum bonuses comes primarily from shifting personnel from extension to reenlistment, and secondarily from increasing the proportion of personnel who choose to stay in the occupation. Our evidence on the cost effectiveness of lump sum bonuses at the second term reenlistment point is inconclusive.
- Reenlistment bonuses, whether lump sum or installment, are effective in increasing the reenlistment rate, decreasing the extension rate, and increasing the retention rate. This pattern implies that higher reenlistment bonuses can increase the expected manyears of active duty service. Lump sum bonuses, which have larger effects on those rates, produce greater expected manyears than installment bonuses.
- An increase in the military/civilian wage index increases both the reenlistment rate and the extension rate, and their sum, the retention rate. The increase in the extension rate, however, is roughly twice as large as the increase in the reenlistment rate. Therefore, although expected manyears clearly increase, a higher fraction of those staying select short commitments. (We have no information on the proportion of extenders who subsequently choose to reenlist.)
- Unemployment effects parallel those of bonuses: higher reenlistment, lower extension, and higher retention rates. As a result, higher bonuses can mitigate the effects of lower

unemployment. Higher military wag[illegible] itself [illegible] unemployment, but unlike *bonuses* they [illegible] increase the proportion of stayers with short commitments.

- Changes in military pay, bonuses, and the unemployment [illegible] that lead to the same change in the *retention rate* can never[illegible] less have different effects *on expected manyears*. For example, given a reenlistment bonus increase and a military pay increase that each produce the same predicted increase in the retention rate for a military occupation, the bonus increase would be associated with a greater increase in expected manyears of service. Models of retention behavior not allowing for these differential effects may give policymakers a misleading impression of the consequences of such changes.
- Estimates of the first and second term effects of a variable were frequently quite similar (Table 4, Sec. VI). This was true for the reenlistment, extension, and retention rate effects of installment bonuses, lump sum bonuses, the military/civilian wage index, and the percent black. However, for the unemployment rate the second term coefficients were larger in absolute value than for the first term, and for the percent without a high school diploma or GED the differences between first and second term were sporadic.
- Controlling for simultaneity bias is essential. Not doing so would have produced a *lower first* term bonus effect on the reenlistment rate and a *negative* second term bonus effect. Our econometric approach seems effective in counteracting the simultaneity bias.
- We identified two sources of simultaneity bias, one associated with persistent yet unobserved factors associated with a military occupation and a second associated with autocorrelation. The first had the greatest effect on our coefficients. The second had little effect on the results because by pooling observations across occupations, we happened to make the average autocorrelation coefficient zero. That neutralized its menace as a source of simultaneity bias. If we had sought to estimate bonus effects for individual occupations or small groups of occupations, autocorrelation might have been a problem.

## BONUSES AS A MANAGEMENT TOOL

Reenlistment bonuses should be viewed as a potent, versatile component of military compensation. Bonuses may be turned on or off rapidly and targeted on critical skills. They increase the retention rate and induce personnel to reenlist rather than extend, thereby increasing expected manyears. The reduction in extensions and increase in reenlistments probably gives personnel managers greater flexibility in planning rotation and new duty assignments, insofar as more personnel are committed to longer terms of service. Bonuses help alleviate transient, unexpected personnel shortages. Less widely recognized, bonuses also provide quasi-permanent pay differentials by occupation that help overcome persistent differences in military/civilian work conditions and earnings opportunities. (Military occupations having a bonus in one period, we observe, tend to have a bonus in the next.) Bonuses can be used to shape the mid-career force—years of service four through twelve—in terms of both size and skill composition. Other studies suggest that such use of bonuses need not affect the number of personnel who continue on toward retirement eligibility, as many of those influenced by bonuses to stay an extra term may leave the service when that obligation is completed.[1] Finally, bonuses have the potential to offset the ill effects on retention of improved national employment conditions. That potential can be achieved, of course, only to the extent that bonus budgets permit sufficient response.

## FUTURE RESEARCH

Our methodology and results help clarify the role of several central policy variables in affecting retention behavior, but several questions remain. With the existing database, future work could analyze *retention behavior by broad occupational group and by service.* This would reveal whether certain occupational groups or services differ substantially from the DoD-level results reported here. Bonus effects by occupational group, term, and service would also be pertinent to assuring the efficient allocation of reenlistment bonus dollars.

An expansion of our database would permit the analysis of further hypotheses. Although extensive tabulations and new variable specifications would be required, Defense Manpower Data Center data could be [illegible] of variables on future *promotion opportunities by occupation* or [illegible] "quality" *composition* of the population of personnel nearing the [illegible] decision point, where the definition of "quality" migh[illegible]

[1] [illegible] and McCall ([illegible]), Warner and Simon ([illegible])

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ratio of the reenlistment rate to the leave rate, therefore the reenlistment logit, which is the log of that ratio.

The idea may be clarified by recalling the expression for the derivative of the reenlistment rate with respect to $x$:

$$\partial r/\partial x = \beta_r r(1 - r) - \beta_e re$$

where $\beta_r$ and $\beta_e$ are the coefficients on $x$ in the reenlistment and extension regressions, respectively. If the latter coefficient were positive and sufficiently large, the partial derivative would become negative. Such complexity can create confusion in interpreting polytomous logit regression results, as the reader is directed from one coefficient to another and asked to keep a mental tally of which effect will prevail.

A preferable alternative, we believe, lies in discussing the hypotheses in terms of the partial derivatives themselves, which by construction show the net effect of the variable on the rate. The derivatives also provide a more natural metric with which to judge the effect of a variable. Another advantage of the partial derivative approach is that the derivatives can be added together. For example, an estimate of the effect of $x$ on the retention rate may be had by adding the partial derivatives of the reenlistment rate and the extension rate with respect to $x$. In our work, we may judge the accuracy of this sum by comparing it with the partial derivative of the retention rate with respect to $x$ computed directly from the retention logit regression.

# Appendix I

## HYPOTHESIS TESTING IN THE LOGIT MODEL

Our hypotheses concern the effect of a variable on a rate, but the dependent variable in the regressions is not a rate but a logit—the log of the ratio of two rates. The dependent variable in the retention regression is the log of the ratio of the retention rate to the leave rate. In the three outcome case—reenlist, extend, or leave—the dependent variable in the reenlistment regression is the log of the ratio of the reenlistment rate to the leave rate, and in the extension regression the dependent variable is the log of the ratio of the extension rate to the leave rate.

In a two outcome (dichotomous) case like retention, hypotheses such as ours can be tested directly with the regression coefficients and their t-statistics. This is because a variable can have a positive effect on a rate if and only if it has a positive effect on its logit. We can see this by examining the partial derivative of the retention rate ($s$) with respect to a variable ($x$):

$$\partial s / \partial x = \beta_x s(1 - s)$$

Here, $\beta_x$ represents the coefficient of $x$ in a regression using the logit of $s$ (the retention rate) as the dependent variable (i.e., $\ln(s/(1 - s))$). The sign of the derivative will be the same as the sign of the coefficient, and the size of the derivative will depend on both $\beta_x$ and $s$. If we evaluate the derivative at a specific value of $s$, treating $s$ as fixed and nonstochastic, the t-statistic of the derivative equals that of the estimated coefficient. For these reasons it is not necessary in the dichotomous logit case to compute the derivative and its t-statistic to test hypotheses stated in terms of the effect of $x$ on $s$.

The situation differs in the polytomous logit case. Even if $x$ has a positive effect on the reenlistment *rate* itself, it need not have a positive effect on the *logit* of the reenlistment rate. Essentially, this ambiguity arises because a change in $x$ can affect both the reenlistment rate and the extension rate. Suppose $x$ increased the reenlistment rate ($r$) and decreased the extension rate ($e$). If the latter effect were stronger, then on net $x$ could increase the leave rate $(1 - r - e)$. Moreover, the leave rate could increase more rapidly than the reenlistment rate. If so, then an increase in $x$ would be associated with a decrease in the

# Appendix H

# CREATION OF MILITARY/CIVILIAN WAGE INDEX

Table H.1

MILITARY/CIVILIAN WAGE INDEX

| | Annual Percent Growth | | Pay Index | | |
|---|---|---|---|---|---|
| Period | Military Base Pay | Avg. Wage Manuf. | Military | Civilian | Military/ Civilian Wage Ratio |
| FY76 II | 2.5 | 4.3 | 102.5 | 104.3 | .976 |
| FY77 I | 2.5 | 4.2 | 105.1 | 108.7 | .959 |
| II | 3.0 | 4.4 | 108.2 | 113.5 | .948 |
| FY78 I | 3.0 | 4.2 | 111.5 | 118.2 | .937 |
| II | 3.2 | 4.5 | 115.0 | 123.5 | .922 |
| FY79 I | 3.2 | 4.2 | 118.7 | 128.7 | .913 |
| II | 4.7 | 3.6 | 124.3 | 133.4 | .912 |
| FY80 I | 4.7 | 4.4 | 130.1 | 139.2 | .915 |
| II | 6.6 | 5.2 | 138.7 | 146.5 | .921 |
| FY81 I | 6.6 | 4.8 | 147.9 | 153.5 | .946 |
| II | 4.6 | 4.0 | 154.7 | 159.6 | .975 |

Military pay increases are for the first term. In FY82 targeted pay increases were given to first and second term personnel. First term pay increased 11.4 percent and second term increased 16.5 percent. Thus in the last six month period of our data the military growth rate is 8.25 for the second term and the base pay index is 157.2. This produces a military/civilian pay ratio of 98.5, which is slightly higher than the 97.5 value shown for the first term.

# Appendix G

## CREATION OF BONUS AMOUNT VALUES

Table G.1 illustrates the steps to construct an occupation's bonus amount for a given period. For each period we obtained base pay from military pay schedules; the monthly base pay values appear in the first two columns. These amounts were deflated by the Consumer Price Index to obtain constant dollar amounts of base pay (1976 dollars). The deflated amounts, shown in columns 3 and 4, were then multiplied by four to reflect the assumed four year reenlistment term. The resulting products (columns 5 and 6) indicate constant dollar bonus amounts per bonus multiple. To compute an occupation's bonus amount for a given term and period, the appropriate value in column 5 or 6 was multiplied by the occupation's prevailing bonus multiple. For example, if an occupation had a first term bonus multiple of three in the second half of FY78, the occupation would be assigned a bonus amount of \$1800 × 3 = \$5400 for that period. Measured in constant dollars, the representative bonus amounts decline by about 10 percent from FY76 to FY81.

Table G.1

CONSTANT DOLLAR BONUS AMOUNTS PER BONUS MULTIPLE

| | Nominal Base Pay | | Deflated Base Pay (CPI index) | | Bonus Amount per Bonus Multiple (1976 dollars) | |
|---|---|---|---|---|---|---|
| Period | First Term (1) | Second Term (2) | First Term (3) | Second Term (4) | First Term (5) | Second Term (6) |
| FY76 II | 487 | 633 | 473 | 614 | 1892 | 2456 |
| FY77 I | 504 | 656 | 475 | 619 | 1900 | 2476 |
| II | 504 | 656 | 458 | 596 | 1832 | 2384 |
| FY78 I | 535 | 696 | 473 | 616 | 1892 | 2464 |
| II | 535 | 696 | 450 | 585 | 1800 | 2340 |
| FY79 I | 565 | 735 | 452 | 588 | 1808 | 2352 |
| II | 565 | 735 | 422 | 549 | 1688 | 2196 |
| FY80 I | 604 | 786 | 422 | 550 | 1688 | 2200 |
| II | 604 | 786 | 403 | 524 | 1612 | 2096 |
| FY81 I | 675 | 878 | 427 | 556 | 1708 | 2224 |
| II | 675 | 878 | 404 | 526 | 1616 | 2104 |

These expressions involving $B_L$ and $B_I$ may be equated and simplified to yield the size of the installment bonus that could be paid per dollar of lump sum bonus. Clearly, recoupment of lump sum bonuses reduces the cost advantage of installment bonuses. For instance, Table F.1 shows that at 8 percent interest and a continuation rate of .95, a $1 lump sum bonus could have funded a $1.12 installment bonus. But in the absence of recoupment the installment bonus would have been $1.20 (Table 8). With full recoupment, cost equivalency is not affected by a change in the proportion of service members continuing from one year to the next.

The figures in Table F.1 ignore the implementation cost of recouping the lump sum bonuses of early leavers. If the implementation cost were *not* negligible, the entries in the table would be higher. Because recoupment would probably not be conducted beyond the point where the cost of implementation exceeded the amount recouped, the entries would be no higher than those in Table 8. Thus Table 8 (no recoupment) and Table F.1 (full recoupment, no cost of recoupment) bracket the size of installment bonus that could be funded at the same cost as a $1 lump sum bonus.

Table F.1

COST EQUIVALENT INSTALLMENT BONUS PER $1 LUMP SUM BONUS, WITH RECOUPMENT
(Assumes a four year term of service)

| | Continuation Rate | |
|---|---|---|
| Interest Rate | .95 | .97 |
| .04 | 1.06 | 1.06 |
| .06 | 1.09 | 1.09 |
| .08 | 1.12 | 1.12 |
| .10 | 1.15 | 1.15 |

# Appendix F

## ALLOWING FOR RECOUPMENT OF LUMP SUM BONUSES

The expected cost of a program of lump sum reenlistment bonuses will be lower if bonus dollars are recouped from personnel leaving before completing their terms of service. Following the presentation in Sec. VII, assume a four year term and a constant continuation rate $p$ from year to year in the term. Also assume that the amount recouped is a simple pro rata share of the bonus, for instance, three-fourths of the bonus for a person completing only one year of the four year term. The size of the lump sum bonus paid to reenlisting personnel is $B_L$.

Expected recoupment per year is determined as follows. A proportion $(1 - p)$ of the reenlisting personnel do not continue after the first year, and $.75B_L$ is recouped from each. The proportion completing the first year and leaving after the second equals $p(1 - p)$, and their recouped amount equals $.50B_L$. Finally, the proportion completing three years then leaving is $p^2(1 - p)$, and their amount is $.25B_L$. The recoupment of these amounts occurs in the years after the reenlistment term begins. However, we assume that the government can effectively borrow from itself to arrange for the intertemporal transfer of funds, hence the present value of the recouped amounts can be obtained by means of discounting at the government's rate of interest.

The cost of a lump sum bonus of amount $B_L$ equals the amount of the bonus itself less the present value of the amount expected to be recouped:

$$B_L[1 - (1 - p).75/(1 + r) - p(1 - p).50/(1 + r)^2$$
$$- p^2(1 - p).25/(1 + r)^3].$$

The size of the installment bonus that can be funded with this cost was derived in Sec. VII:

$$.25B_I[1 + p/(1 + r) + p^2/(1 + r)^2 + p^3/(1 + r)^3].$$

# Appendix E

# REENLISTMENT, EXTENSION, RETENTION REGRESSION RESULTS: METHOD II

Table E.1

WEIGHTED, CENTERED DATA; SEEMINGLY UNRELATED REGRESSIONS
(t-statistic)

| Variable | First Term Reenlist | First Term Extend | First Term Retention | Second Term Reenlist | Second Term Extend | Second Term Retention |
|---|---|---|---|---|---|---|
| Has a bonus | .0098<br>(0.31) | .0189<br>(0.40) | −0.017<br>(−0.59) | .1968[a]<br>(4.53) | −0.033<br>(−0.52) | .1429[a]<br>(3.58) |
| Has a bonus after 4/79[b] | .0921[a]<br>(3.03) | .0528<br>(1.07) | 0.087[a]<br>(2.96) | .0117<br>(0.26) | .1230<br>(1.82) | .0490<br>(1.17) |
| Bonus amount ($1000) | .0759[a]<br>(11.34) | −.0275[a]<br>(−2.64) | .0490[a]<br>(7.63) | .0323[a]<br>(4.00) | −.0116<br>(−0.93) | .0180[c]<br>(2.45) |
| Bonus amount after 4/79 ($1000)[b] | .0299[a]<br>(4.57) | −.0173<br>(−1.60) | .0146[c]<br>(2.28) | −.0077<br>(−1.06) | −.0380[a]<br>(−3.26) | −.0145[c]<br>(−2.20) |
| Military/civilian wage index | .0266[a]<br>(6.01) | .0855[a]<br>(14.14) | .0466[a]<br>(11.33) | .0245[a]<br>(5.04) | .0757[a]<br>(11.48) | .0372[a]<br>(8.30) |
| Unemployment rate in previous period | .0875[a]<br>(7.55) | −0.099[a]<br>(−6.25) | .0284[a]<br>(2.64) | .1483[a]<br>(10.89) | −.0677[a]<br>(−3.63) | .1027[a]<br>(8.17) |
| Percent with no HS diploma | .0001<br>(0.11) | −.0049[a]<br>(−5.47) | −.0020[a]<br>(−3.26) | −.0074[a]<br>(−5.36) | .0092[a]<br>(5.31) | .0029[c]<br>(−2.32) |
| Black | .0109[a]<br>(16.01) | .0083[a]<br>(8.27) | .0096[a]<br>(14.71) | .0050[a]<br>(7.45) | .0086[a]<br>(9.42) | .0061[a]<br>(9.66) |

[a]Significant at 99 percent confidence level (critical t = 2.58).

[b]Coefficient represents the difference between the installment period (prior to 4/79) and the lump sum period (4/79–12/81). For lump sum effect, the coefficient is added to that for pre 4/79 period.

[c]Significant at the 95 percent confidence level (critical t = 1.96).

# Appendix D

# REENLISTMENT, EXTENSION, RETENTION REGRESSION RESULTS: METHOD I

Table D.1

WEIGHTED DATA; SEEMINGLY UNRELATED REGRESSIONS
(t-statistic)

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| Variable | Reenlist | Extend | Retention | Reenlist | Extend | Retention |
| Has a bonus | .1158[a] | -.3569[a] | -.0335 | .2566[a] | -.0052 | .2072[a] |
| | (3.46) | (-6.78) | (-1.13) | (5.51) | (-0.07) | (4.68) |
| Has a bonus after 4/79[c] | -.0996[b] | -.0040 | -.0693 | .0730 | .2040[b] | .0998 |
| | (-2.45) | (-0.06) | (-1.93) | (1.29) | (2.16) | (1.85) |
| Bonus amount ($1000) | .0173[b] | -.0203 | .0077 | -.1085[a] | -.1144[a] | -.1082[a] |
| | (2.55) | (-1.86) | (1.28) | (-15.2) | (-8.7) | (-16.1) |
| Bonus amount after 4/79 ($1000)[c] | .0560[a] | .0150 | .0413[a] | .0324[a] | .0197 | .0263[a] |
| | (6.08) | (1.03) | (5.06) | (3.68) | (1.25) | (3.17) |
| Military/civilian wage index | -4.121[a] | -3.609[a] | -3.183[a] | -2.7997[a] | -3.0212[a] | -2.281[a] |
| | (-58.5) | (-39.8) | (-53.9) | (-33.9) | (-23.9) | (-29.0) |
| Unemployment rate in previous period | .2485[a] | .1771[a] | .2097[a] | .2804[a] | .1555[a] | .2437[a] |
| | (27.9) | (15.4) | (28.0) | (26.5) | (9.58) | (24.16) |
| Percent with no HS diploma | -.0067[a] | -.0062[a] | -.0080[a] | .0059[a] | .0339[a] | .0052[a] |
| | (-11.5) | (-9.01) | (-17.2) | (-3.54) | (22.8) | (5.03) |
| Percent black | .0176[a] | .0000 | .01246[a] | .0152[a] | .0026 | .0128[a] |
| | (42.1) | (-0.05) | (34.12) | (27.4) | (2.44) | (24.29) |

[a]Significant at the 99 percent confidence level (critical t = 2.58).
[b]Significant at the 95 percent confidence level (critical t = 1.96).
[c]Coefficient represents the difference between the installment period (prior to 4/79) and the lump sum period (4/79–12/81). For lump sum effect, the coefficient is added to that for pre 4/79 period.

In our preliminary analysis, we ran a SUR model allowing for covariance between the error terms in the reenlistment and extension logit regressions. We found the covariance to be low; the equations were nearly independent. Further, even if the covariance had not appeared to be negligible, it is evident from the expressions for the variances of the derivatives that the covariance of the parameter estimates is multiplied by terms that are quite small; consequently, the effect of the covariance is diminished. As a result of this fact and the low estimated error covariance between the logit regressions, we decided to assume that the covariances between the estimated parameters of the logit regression were zero. This decision led to the following expressions for the variances of the partial derivatives.

$$\sigma^2_{\frac{\partial r}{\partial x}} = r^2(1 - r)^2\sigma^2_{\beta_r} + r^2e^2\sigma^2_{\beta_e}$$

$$\sigma^2_{\frac{\partial e}{\partial x}} = e^2(1 - e)^2\sigma^2_{\beta_e} + r^2e^2\sigma^2_{\beta_r}$$

In evaluating these expressions, we treat the reenlistment and extension rates as nonstochastic and set them at the average values for the first term and the second term, respectively. From these estimated variances the standard errors are computed directly by taking the square root, and the t-statistic of the derivative is determined by forming the ratio of the value of the derivative to its estimated standard error.

Although this approach produces only approximations to the true t-statistics, we believe the approximations to be accurate in the case of our research.

# Appendix C

## APPROXIMATE T-STATISTICS FOR LOGIT PARTIAL DERIVATIVES

Our approach to estimating the polytomous logit model of the reenlistment rate and the extension rate uses Berkson's approximation, wherein each logit may be written as a linear equation with a normally distributed error term. This results in a system of $n - 1$ equations for a polytomous choice involving $n$ outcomes. In general the error terms may be correlated across equations. If so, the standard error of an expression involving the estimated parameters from, say, both the reenlistment logit regression and the extension logit regression would have a standard error involving a possible covariance between the two parameters.

The formulas estimating the first derivatives of the reenlistment rate and the extension rate with respect to an explanatory variable $x$ are:

$$\frac{\partial r}{\partial X} = \hat{\beta}_r r(1 - r) - \hat{\beta}_e re$$

$$\frac{\partial e}{\partial X} = \hat{\beta}_e e(1 - e) - \hat{\beta}_r re$$

Each derivative contains a parameter from both logit regressions, and the estimates of the parameters come from the logit regressions. The variances of the derivatives may be written as

$$\sigma^2_{\frac{\partial r}{\partial x}} = r^2(1 - r)^2\sigma^2_{\beta_r} + r^2e^2\sigma^2_{\beta_e} - 2(r^2(1 - r)e)\sigma^2_{\beta_r\beta_e}$$

$$\sigma^2_{\frac{\partial e}{\partial x}} = e^2(1 - e)^2\sigma^2_{\beta_e} + r^2e^2\sigma^2_{\beta_r} - 2(e^2(1 - e)r)\sigma^2_{\beta_e\beta_r}$$

Note that these expressions involve the standard error for each estimated parameter as well as a term for the covariance between the estimates.

# Appendix B

# REENLISTMENT, EXTENSION, RETENTION REGRESSION RESULTS: FINAL METHOD

Table B.1

WEIGHTED, CENTERED, QUASI-DIFFERENCED DATA, GENERALIZED LEAST SQUARES
(t-statistic)

| | First Term | | | Second Term | | |
|---|---|---|---|---|---|---|
| Variable | Reenlist | Extend | Retention | Reenlist | Extend | Retention |
| Has a bonus | .0262<br>(0.86) | .0400<br>(0.84) | -.0136<br>(-0.54) | .1877[a]<br>(4.87) | -.0179<br>(-0.29) | .0943[a]<br>(3.00) |
| Has a bonus after 4/79[c] | .0536<br>(1.92) | .0026<br>(0.05) | .0303<br>(1.20) | .0265<br>(0.66) | -.0140<br>(-0.21) | .0571<br>(1.70) |
| Bonus amount ($1000) | .0737[a]<br>(11.92) | -.0280[a]<br>(-2.61) | .0490[a]<br>(9.09) | .0483[a]<br>(6.85) | -.0188<br>(-1.57) | .0310[a]<br>(5.20) |
| Bonus amount after 4/79 ($1000)[c] | .0263[a]<br>(4.23) | -.0265[a]<br>(-2.34) | .0100<br>(1.77) | -.0073<br>(-1.23) | -.0297[a]<br>(-2.91) | -.0141[a]<br>(-2.77) |
| Military/civilian wage index | .0308[a]<br>(5.73) | .0827[a]<br>(11.8) | .0482[a]<br>(11.1) | .0205[a]<br>(3.07) | .0688[a]<br>(7.72) | .0318[a]<br>(5.66) |
| Unemployment rate in previous period | .1127[a]<br>(8.33) | -.0448[b]<br>(-2.56) | .0625[a]<br>(5.77) | .1408[a]<br>(7.65) | -.000646<br>(-0.03) | .1196[a]<br>(7.70) |
| Percent with no HS diploma | .000977<br>(1.37) | -.00339[a]<br>(-3.84) | -.000764<br>(-1.38) | -.00531[a]<br>(-3.68) | .00601[a]<br>(3.25) | -.00232[b]<br>(-1.98) |
| Percent black | .00977[a]<br>(13.9) | .00628[a]<br>(6.22) | .00851[a]<br>(14.3) | .00566[a]<br>(8.33) | .00557[a]<br>(5.97) | .00603[a]<br>(10.8) |

[a] Significant at the 99 percent confidence level (critical t = 2.58).
[b] Significant at the 95 percent confidence level (critical t = 1.96).
[c] Coefficient represents the difference between the installment period (prior to 4/79) and the lump sum period (4/79-12/81). For lump sum effect, the coefficient is added to that for pre 4/79 period.

# Appendix A

## MEANS AND STANDARD DEVIATIONS OF VARIABLES

| Variable | Unweighted | | Weighted[a] | |
|---|---|---|---|---|
| | First Term | Second Term | First Term | Second Term |
| Reenlistment Rate | 13.7 (11.6) | 35.3 (19.1) | 13.9 (6.8) | 35.4 (14.0) |
| Extension Rate | 8.3 (9.2) | 11.4 (13.2) | 7.0 (4.8) | 10.1 (8.7) |
| Retention Rate | 21.9 (14.1) | 46.6 (20.2) | 21.0 (7.6) | 45.5 (15.3) |
| Has a bonus | 34.6 (47.5) | 22.6 (41.9) | 36.5 (48.2) | 23.9 (42.7) |
| Bonus amount in 1976 dollars | 4295 (2198) | 4679 (3064) | 3732 (2310) | 6112 (4426) |
| Military/civilian wage index | 93.86 (2.29) | 93.95 (2.45) | 93.78 (2.34) | 93.89 (2.47) |
| Unemployment rate in previous period | 7.10 (0.88) | 7.10 (0.88) | 7.09 (0.88) | 7.07 (0.86) |
| Percent without HS diploma or GED | 11.8 (14.0) | 5.4 (9.1) | 17.1 (15.3) | 6.1 (14.7) |
| Percent black | 17.6 (15.7) | 18.1 (16.5) | 23.3 (12.3) | 21.2 (7.0) |

[a]Weighted by the number of personnel in each occupation and period.

include entry-level information as well as promotion history; on the proportion of personnel who received a *prior bonus*; and on the likelihood of receiving a *future bonus* at the next reenlistment point. DMDC data can also help *adjust the extension rate* for the fraction of extenders in the process of retraining.

DMDC personnel files could track the extent of *occupational migration* in the services, much of which apparently occurs at the end of the first term and could, on net, be beneficial to maintaining overall retention. Analysis would be required to determine whether reenlistment bonuses affect occupational migration, or whether such flows are largely governed by service policies regarding the movement among occupations that are under, at, or over required strength. Finally, improved data on *reenlistment eligibility* would be highly desirable. The development of a specific chronology by service and occupation of reenlistment eligibility policy, and its implementation, would be of great value in policy analysis.

The policy payoff to further work on the existing and enhanced databases comes in the form of improved knowledge of the retention effects of pay, unemployment, bonus, promotion, and eligibility variables. Such knowledge can be applied in several ways. For instance, our current model can be used to forecast first and second term retention behavior.[2] The model could also simulate how the future path of retention rates would change under different reenlistment bonus scenarios, characterized in terms of bonus budget, coverage, and payment amounts. As suggested, the addition of variables concerning promotion and eligibility would strengthen the forecasting capability of the model.

In another direction, retention analyses provide needed information on whether planned changes in military pay and bonuses are adequate to meet manning requirements. Moreover, although most analyses of retention behavior have dealt with personnel *per se*, in our opinion it would be valuable to determine whether current pay, bonus, promotion, and eligibility policies tend to favor the retention of personnel who appear to have been better performers, or who appear to have the potential to be high future performers. Finally, the gradual accumulation of information about bonus effects by occupational group is useful for [illegible] an efficient allocation of the reenlistment bonus budget.

[2] [illegible] forecasts based on the model have been [illegible] trend [illegible] structure of the enlisted force (Hosek, Fernandez, [illegible])

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